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American Journal of
**PUBLIC
HEALTH**

A PUBLICATION OF
AMERICAN PUBLIC HEALTH ASSOCIATION

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Impact of COVID-19 on Primary Health Care Services

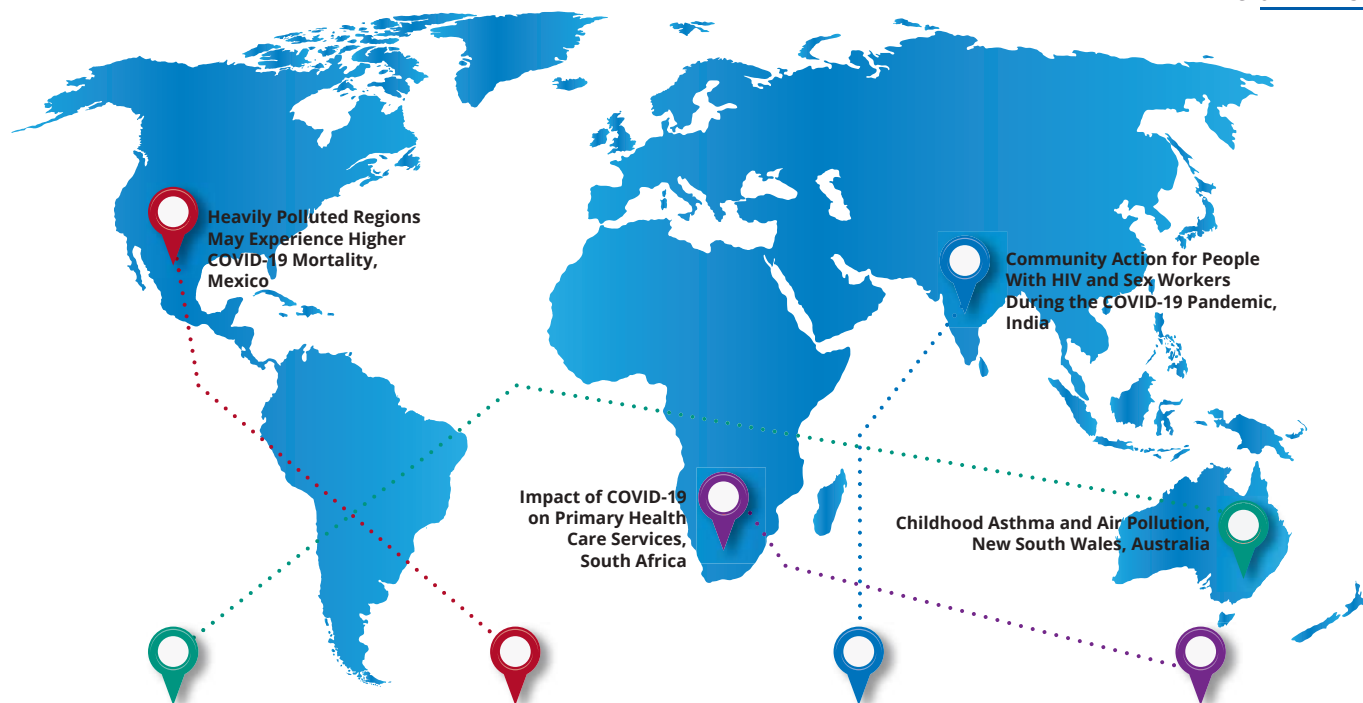
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Prepared by Stephen A. Lewandowski, Rebekah C. Hughes, Megan E. Marziali, and Luis E. Segura. Columbia University, New York, NY. Correspondence should be sent to the AJPH Global News Team at les2196@cumc.columbia.edu.

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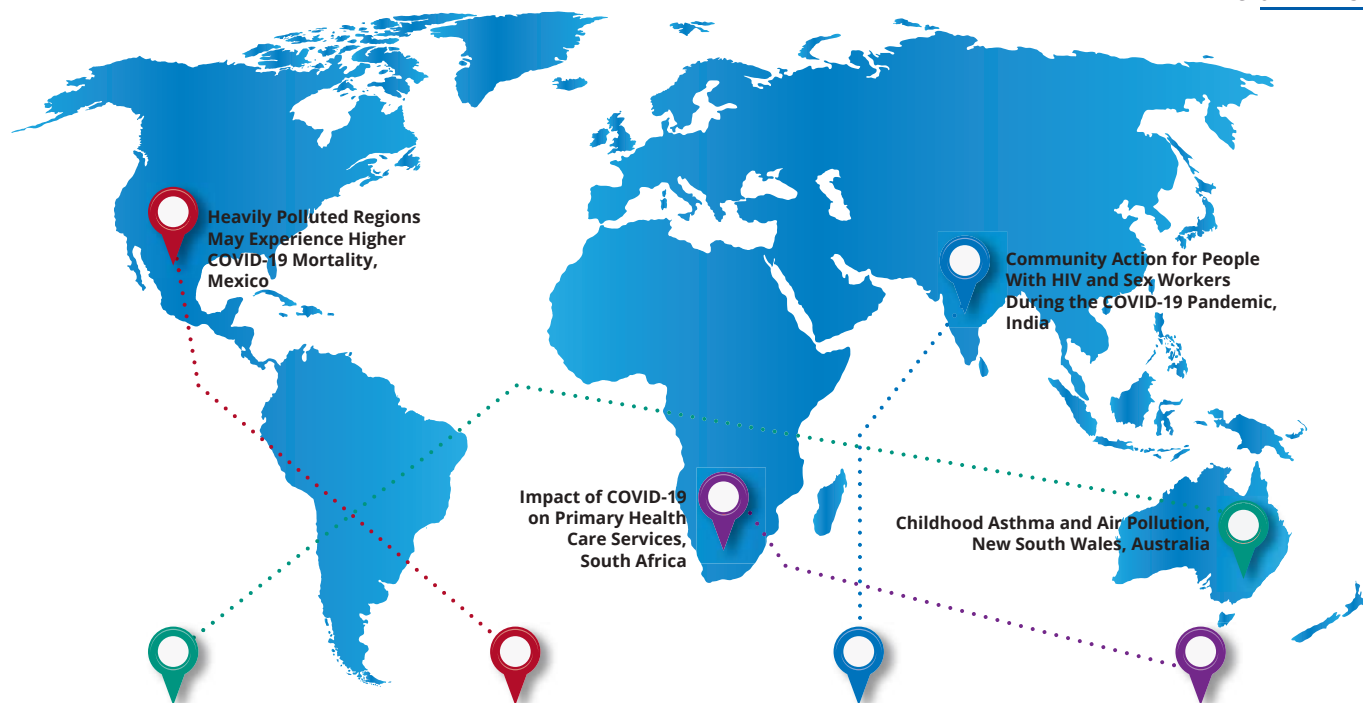
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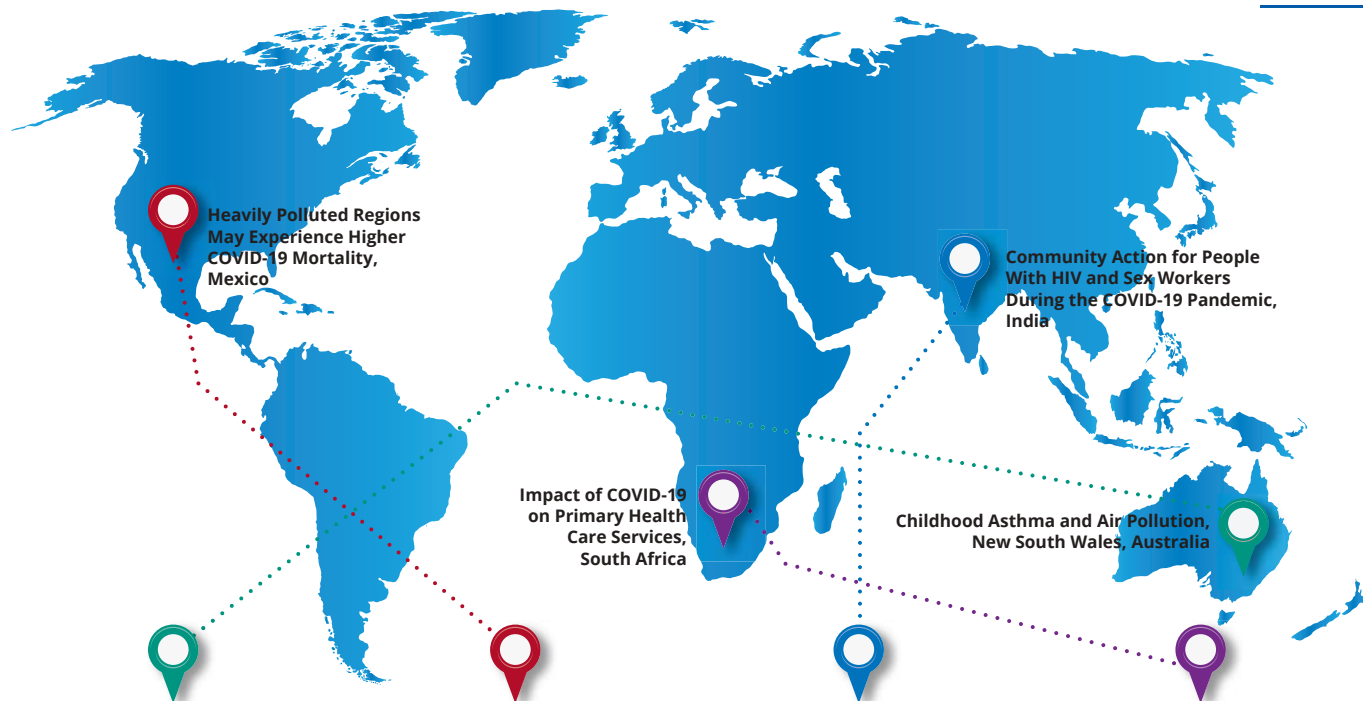
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COVER: The effects of COVID-19 are far more vast than a positive diagnosis. The related recession and the negative effect on the economy is responsible for an important amount of excess deaths (see Brenner et al., p. 1950). People, especially communities of color, are drowning in a sea of adjacent issues exacerbating and adding to the stress of the pandemic.

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Choosing to Be Both: Perspectives From a Researcher About Early Motherhood



Over the past year, I have had the privilege of adding a number of professional achievements to my curriculum vitae, including becoming the student editor for *AJPH*. There is no place on my curriculum vitae, however, for the most important and life-changing new role I have taken on this year: mother.

I approached the process of becoming a mother in the same way that I would approach any new “project” I conjured up as a doctoral student—I naturally fell into the researcher mentality whereby I began analyzing, studying, asking questions, and preparing for everything I could possibly think of regarding babies and becoming a new mother. I conducted informal interviews with friends and family members, created excel files to assess daycare centers, reviewed the pros and cons of different strollers, attended classes on breast feeding, read scientific literature about infant sleep, and made an infographic of my birth plan for hospital staff. Even after my son arrived, I continued to use this scientific process—convinced that I could “research” my way through anything that arose as part of new motherhood. Phone apps made it possible to collect data about my son’s feedings, bowel movements, nap times, and growth, just to name a few. With these data I developed hypotheses about almost everything. Most notably, I focused on what might be causing baby to be colicky at night and worked to ensure that all possible confounding variables were controlled for to guarantee that the baby would sleep through the night.

Turns out, approaching motherhood from a purely scientific mindset was unsustainable. The approach I had chosen (one in which I was in complete control) had failed me. The rigid expectations of what my motherhood experience *should* be like—from labor and delivery to how quickly my body and mind would “bounce back”—were dismantled. I sought to maintain control of the “researcher” mentality whereby I was able to identify causal

pathways and test interventions to manipulate the predictors related to my “participant’s” (son’s) outcomes of interest. Slowly, I realized that it was unachievable and downright miserable to try to hold on to control in this way. The reality was that motherhood, especially during a pandemic, required adaptation or would lead very quickly to burnout. And so, I gradually shifted from the comfortable role I had always held as an outside observer and inquisitive problem solver to an active participant of a live and ever-changing experiment. The reality was that in this motherhood role, I had to be “both . . . and”: *both* the prepared and thoughtful investigator *and* the present and active participant.

So, how has this new role, which is completely unaccounted for as part of my professional identity but simultaneously the most life-changing and long-term position I will ever hold, changed my approach to public health and research? At the core, I think it has provided an important gateway for me to be open, honest, and vulnerable, even in professional settings, which in turn gives others permission to do the same. It also encourages me to advocate institutionalized systems that support and nurture the balance between the two roles successfully. The messy nature of early motherhood has helped me appreciate the complex aspects of research, challenging me to maintain high standards in my research while also becoming more accepting of the uncertainty that comes with studying human behaviors. I have learned to ask new and better questions, to think about things from both the scientific and the practical side. I now understand that it is possible to be ambitious and flexible, determined and gentle, dedicated to my career and my family. In short, a mother *and* a researcher. **AJPH**

Caitlin G. Allen, PhD, MPH
Postdoctoral Fellow
Medical University of South Carolina,
Charleston

DOI: <https://doi.org/10.2105/AJPH.2021.306528>

10 Years Ago

Impact of Business Cycles on US Suicide Rates

Our findings suggest the importance of population-level suicide-prevention strategies, particularly during recessions. For example, during recessions importance must be placed on providing social support and counseling services to those who lose jobs or home; promoting individual, family, and community connectedness; and providing adequate resources to crisis call centers and other community services. . . . We found that people in prime working ages (25-64 years) were more vulnerable to recession than were others. This may be partly explained by the fact that many of those people were breadwinners in their homes, and their jobs supported mortgage payments, health insurance, children’s education, and other expenses. . . . The multifaceted nature of suicide indicates the need to develop prevention efforts that use multiple settings where vulnerable people and individuals may be found.

From *AJPH*, June 2011, pp. 1143–1145 *passim*

50 Years Ago

Economic Changes and Heart Disease Mortality

Among the major social stresses that may be involved in heart disease incidence or mortality are those which originate in adverse changes in the economic status of individuals. . . . Periods of instability, particularly periods of economic recession, force a sizable portion of the population either out of the economy. . . . or into a situation of decreased income. . . . It is . . . a reasonable speculation that heart disease mortality might increase during an economic downturn as a result of the decreased availability, in economic terms, of medical care services. . . . The findings of the present study clearly indicate that economic downturns are associated with increased mortality from heart disease and that, conversely, heart disease mortality decreases during economic upturns.

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Warning: Don't Let the Beverage Industry Harm Your Kids

Lori Dorfman, DrPH, MPH

ABOUT THE AUTHOR

Lori Dorfman is with the Division of Community Health Sciences, School of Public Health, University of California, Berkeley, and the Berkeley Media Studies Group, Public Health Institute, Berkeley, CA.

 See also Krieger et al., p. 1997.

This issue of *AJPH* brings some good news for everyone concerned about childhood nutrition: in a rigorous, randomized controlled trial—the first study of its kind to my knowledge—Krieger et al. (p. 1997) found that countermarketing on social media could reduce purchases and consumption of sugary fruit drinks and increase consumption of water among Latino/a/x¹ parents and their children. The key to understanding the significance of this study is in the definition of “countermarketing,” a technical term that directs audiences’ attention to the behavior of industry.

DENORMALIZING INDUSTRY PRACTICES

Unlike general social marketing or other communications campaigns, countermarketing is designed “to reduce the demand for unhealthy products by exposing the motives of their producers and portraying their marketing activities as outside the boundaries of civilized corporate behavior.”^{2(p120)} The spectrum of health communications can stretch from inspiring individual behavior change on one end to campaigns that influence

policy, systems, and environmental change on the other; countermarketing sits at the systems end of the spectrum, with the intention of shining the light on institutions with the power to change environments.³ At scale, countermarketing improves health by denormalizing the marketing practices—the normal business practices—of companies producing products that harm health.²

The messages Krieger et al. tested put the beverage industry’s manipulative marketing front and center. For example, one of their messages shows a girl with severe tooth decay holding a fruit drink pouch bearing an “all-natural” claim and the text “Just because a label states ‘all-natural’ doesn’t make a fruit drink healthy. Don’t let the beverage industry harm your kids.” (I encourage *AJPH* readers to view the images in the supplemental material to feel their, well, punch.) The study demonstrated the strength of the messages with statistically significant reductions in parents selecting fruit drinks in the virtual store and in their reported consumption in real life later.

As Krieger et al. note, tobacco control has had celebrated success with countermarketing. Indeed, the study’s

message is reminiscent of a countermarketing message from California’s storied campaign Warning: The Tobacco Industry Is Not Your Friend, a campaign credited with tripling the decline in smoking in California.⁴ But such campaigns are expensive. California’s multimillion dollar tobacco education campaign was funded by a statewide tobacco tax, and a similarly robust early effort, the original “truth” campaign in Florida, was funded with \$200 million from Florida’s \$11.3 billion settlement with the tobacco industry.⁵ Krieger et al. suggest that their findings will help local organizations without such deep pockets reach parents with successful countermarketing because the messages were tested on social media, in this case, Facebook.

QUESTIONS OF HEALTH AND RACIAL EQUITY

Herein lies an irony: is Facebook the solution or the problem? Krieger et al. point out the affordability of social media campaigns, but reach is still an issue. Posts from alcohol, tobacco, and food companies often generate millions of views, whereas community-organized grassroots health-focused social media might generate views that number in the hundreds.⁶ Getting to scale, even on social media, can be expensive.

The social media platforms themselves, including Facebook, are now the largest junk food marketers. Google and Facebook alone account for the majority (nearly 60%) of digital ad spending worldwide,⁷ and social media is where food and beverage companies go to reach children.⁸ Facebook has worked with Coca-Cola, PepsiCo, Unilever, Nestlé, and many other companies to enable sophisticated marketing

across its platform.⁸ Consider just one example: in 2019, Pepsi partnered with Facebook's Instagram to produce 230 million bottles of soda imprinted with mobile phone codes that triggered "full-screen immersion" augmented reality effects to stimulate purchases of 20-ounce bottles of Pepsi.⁸ Even with the low entry fees to social media, it will be hard for public health to compete with immersive, enticing, and ever-present digital campaigns.

Sugary beverage marketing is a health equity issue because the food, beverage, and digital marketing industries hold the power over what children see. It is also a racial equity issue because children of color are heavily targeted with marketing for sugary beverages and other junk food.⁹ Although parents control the products they serve their children, it is not fair to hold parents solely responsible for the effects of marketing they do not control. These power imbalances force us to turn our attention not just to the decisions individual parents make about what they and their families consume but also to the policies that shape the environment in which those decisions are made.

Countermarketing can help when the messages are part of policy campaigns to rein in industry behavior. Recent research has demonstrated, for example, that messages emphasizing the intersections among industry behavior, parental decisions, and community efforts may be particularly effective in moving diverse constituents to support policy such as marketing restrictions, and communities of color in particular may be more attuned, perhaps through lived experience with aggressive target marketing, to the value of policy in shaping industry behavior.¹⁰ Current legislative and legal antitrust actions

against tech giants, including Facebook, are an opportunity to restrict marketing practices across digital platforms, including marketing for sugary drinks.⁸

CONCLUSIONS

In the early 1990s, the task of the tobacco control movement was "to publicly identify the tobacco cartel as the enemy, and to fight to dislodge it from its positions of power."¹¹(p35) Countermarketing was an important tool in tobacco control. Krieger et al. have shown that we can wield the same tool to improve childhood nutrition by making visible the actions of food and beverage marketers targeting children. Let's put this good news to use by placing responsibility at the feet of the platforms and companies profiting from directly and incessantly targeting children with marketing the food and drink they should avoid. **AJPH**

CORRESPONDENCE

Correspondence should be sent to Lori Dorfman, Berkeley Media Studies Group, 2130 Center St, Suite 302, Berkeley, CA 94704 (e-mail: dorfman@bmsg.org). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONFLICTS OF INTEREST

The author has no conflicts of interest to declare.

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Gun Violence Prevention: A Public Health Approach

Edited By: Linda C. Degutis, DrPH, MSN, and Howard R. Spivak, MD



Gun Violence Prevention: A Public Health Approach acknowledges that guns are a part of the environment and culture. This book focuses on how to make society safer, not how to eliminate guns. Using the conceptual model for injury prevention, the book explores

the factors contributing to gun violence and considers risk and protective factors in developing strategies to prevent gun violence and decrease its toll. It guides you with science and policy that make communities safer.

2021, SOFTCOVER, 230 PAGES, 9780875533117



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New Evidence on Sweetened Beverage Taxes Points the Way for Future Policy and Research

Melissa Knox, PhD

ABOUT THE AUTHOR

Melissa Knox is with the Department of Economics, University of Washington, Seattle.

 See also Chrisinger, p. 1986.

In the six years since Berkeley, California, enacted its sugar-sweetened beverage (SSB) tax, seven other US cities have followed suit, recognizing these taxes as feasible policy both for addressing public health challenges related to excess sugar consumption and for increasing revenues to invest in food access and other social determinants of health. The proposed mechanism behind the taxes is simple: as excise taxes, they lead to increased prices for taxed beverages, and higher prices cause consumers to reduce their SSB purchases in favor of healthier beverages. Ultimately, it is hoped, taxes will reduce added sugar intake, thus reducing obesity and improving health, while revenues from remaining sweetened beverage sales are invested back into community health efforts.

Although there is consistent evidence that the prices of taxed beverages increase after SSB taxes are imposed, tax effects on household SSB purchases and intake are less clear.¹⁻³ Studies using retail scanner data consistently find that purchases of taxed

beverages fall at retailers located in taxed cities.^{4,5} But some of these studies have found contemporaneous increases in purchases of SSBs at retailers located outside the city's borders,⁵ suggesting that residents may simply be avoiding taxes without changing their intake.

Along with doubts about household-level responses to SSB taxes come questions about tax equity. As with sales taxes, SSB taxes may impose higher burdens on low-income households than they do on higher-income households.⁶ Unlike a sales tax, however, low-income households can reduce their SSB tax burden by reducing purchases of taxed products or by traveling outside the jurisdiction to purchase SSBs. Because SSB taxes in the United States are city-level taxes, avoiding taxes by crossing borders to shop is relatively easy, but evidence of cross-border shopping by income group is thin.

Given the policy implications of inequitable tax burdens, and the tendency of the beverage industry to use equity arguments to speak out against SSB

taxes, it is important to fully understand household responses to these taxes, both overall and by income.⁷ There is little evidence on this topic so far, mainly because of the need for researchers to collect primary survey data to investigate household-level behavior in a single city.^{8,9}

NEW DATA SHINES A NEW LIGHT ON THE ISSUE

In this issue of *AJPH*, Chrisinger (p. 1986) uses a unique data source to improve our knowledge about low-income household responses to SSB taxes. Using administrative data on county-level redemptions of Supplemental Nutrition Assistance Program (SNAP) benefits, Chrisinger investigated SNAP redemption patterns in Philadelphia, Pennsylvania, and its neighbor counties before and after the enactment of Philadelphia's 2017 sweetened beverage tax. Chrisinger found that total SNAP redemption and SNAP redemption per participant increased in several of Philadelphia's neighbor counties after Philadelphia implemented its tax, even though SNAP redemptions fell across the United States during the same period.

Chrisinger used the method of synthetic controls to estimate the proportion of these post-tax increases that can be attributed to the tax. He found that the tax was associated with up to \$6 million in increased redemptions in Montgomery, Bucks, and Delaware counties, the three Pennsylvania counties bordering Philadelphia. He was not able to estimate the decline in total redemptions in Philadelphia because of the tax but found that per participant SNAP redemption fell there by about \$10 per month. At the same time, per

participant SNAP redemptions increased by about \$35 and \$20 per month in Montgomery and Delaware counties, respectively.

These findings provide new and valuable evidence regarding the cross-border shopping habits of low-income Philadelphia residents. However, the study had some limitations. Chrisinger's inability to estimate the post-tax decline in SNAP redemptions in Philadelphia makes it difficult to assess the importance of the cross-border increases he measured. He did find that per participant SNAP redemptions increased more in Philadelphia's neighbors than they declined in Philadelphia. However, with different numbers of participants in each county, it is again difficult to determine the relevance of the relative quantities.

It is also possible that Chrisinger's findings are not generalizable to other US sweetened beverage taxes. Philadelphia is unique in taxing beverages sweetened with noncaloric sweeteners along with SSBs. Nearly twice as many ounces of taxed beverages were purchased per capita in Philadelphia in 2020 than in Seattle, Washington, or San Francisco, California, two other large cities with SSB taxes, suggesting a much larger tax burden in Philadelphia.¹⁰

Finally, although the synthetic control method is the standard for conducting policy evaluation with a small number of treated units, it suffers from the usual limitations of causal inference with observational data.¹¹ So although Chrisinger was able to conclude that the patterns associated with Philadelphia's sweetened beverage tax are not attributable to random chance, it is impossible to say that the changes were caused by the tax and not by co-occurring but unobserved economic or behavioral factors.

Despite these concerns, Chrisinger's results offer important new insights into behavioral responses to SSB taxes. Future researchers can and should use these data to replicate the work for the other seven SSB taxed cities in the United States. Doing so will produce more robust and more generalizable estimates of the impact of SSB taxes on food purchasing behavior of the cities' low-income residents.

IMPLICATIONS FOR FUTURE POLICIES

Chrisinger's findings supply the first evidence, to my knowledge, that post-SSB tax cross-border shopping trends can be driven by low-income as well as high-income consumers, although it is difficult to determine the extent to which low-income families are moving their shopping trips outside the city. Leaving the city to purchase groceries imposes time, transportation, and convenience costs on consumers, though, and these costs are likely to be significant for low-income purchasers. That cross-border shopping is costly can be seen in previous estimates that sweetened beverage purchases outside Philadelphia only account for about 25% of the sweetened beverage sales lost by Philadelphia retailers.⁵ Still, greater geographic spread of SSB taxes, particularly the introduction of statewide SSB taxes, is likely to make tax avoidance more costly and further reduce household SSB purchases.

These results also raise new questions about the broader impacts of sweetened beverage tax policy and whether cities should act to reduce these effects. Chrisinger suggests that some SNAP beneficiaries are moving entire shopping trips outside Philadelphia city limits because of the tax. This

claim is worthy of further investigation into whether SSB taxes cause negative spillovers to the revenues of food retailers. If these effects are found, Philadelphia could consider using sweetened beverage tax revenues to subsidize healthy food purchases made with SNAP dollars to promote healthy diets while reducing incentives to leave the city to purchase groceries. Seattle has implemented a similar policy, using part of its sweetened beverage revenues to fund the Fresh Bucks Match program, which is available at supermarkets and neighborhood grocers. Another way for Philadelphia to increase the effectiveness of its tax and discourage cross-border shopping is to invest in public health information campaigns to reduce demand for SSBs.¹⁰ In the end, pairing price increases with these types of mitigation strategies may do more than a price increase alone to permanently change dietary patterns and promote both equity and public health in the long run. *AJPH*

CORRESPONDENCE

Correspondence should be sent to Melissa Knox, UW Economics, Box 353330, Savery Hall 305, University of Washington, Seattle, WA 98195-3330 (e-mail: knoxm@uw.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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AJPH Call for Papers

Ubiquitous Lead

The *American Journal of Public Health (AJPH)*, in collaboration with the Centers for Disease Control and Prevention (CDC), intends to publish a supplemental issue on the topic of lead hazards, prevention–mitigation programs, and emerging sources of exposure. The supplement will address and contribute to the comprehensive understanding of currently known and emerging hazardous sources of lead exposure related to global trade, climate change, and infrastructure renewal. Original social, policy, research and evaluation articles, and perspectives are invited. Topics of interest include but are not limited to evidence-based promising practices that strengthen efforts to identify, measure, and mitigate lead exposure in communities; analyses of data demonstrating geographic distribution of lead exposures and associated social vulnerabilities; social determinants of lead exposure risk and consequences; blood lead level testing in affected communities; building capacity for prevention and working with agency partners; emerging sources of lead exposure; implications of COVID-19 on lead poisoning prevention and surveillance; lead poisoning as a major international health crisis; and data demonstrating the impact of lead on children and adults. The full Call for Papers is available at <https://ajph.aphapublications.org/callforpapers>.

Scholars in academia, historians, public health departments, the medical professions, health educators and evaluators, community and faith-based organizations, and governmental agencies are invited to submit manuscripts related to lead exposure prevention and mitigation. Potential authors should visit the *AJPH* website (<https://www.ajph.org>) to review the Instructions for Authors and specific guidelines for the various types of manuscripts. Importantly, submissions must include a cover letter formatted as requested in the Instructions for Authors. In all manuscripts, the number of words, references and tables/figures must correspond to a specific *AJPH* article format. All manuscripts will undergo standard peer review by the *AJPH* editors and peer referees as defined by *AJPH* policy. Manuscripts must be submitted to *AJPH* by **January 30, 2022**, and can be submitted at <https://www.editorialmanager.com/ajph>. For additional information about this supplement, contact: T. LeBlanc at tbleblanc@cdc.gov.

Guest Editors: Tanya Telfair LeBlanc, PhD, MS; Erik Svendsen, PhD; and Paul Allwood, PhD, Centers for Disease Control and Prevention, Atlanta, GA

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New Directions for Public Health Research on the Health and Health Care of Undocumented Immigrants

Alexander N. Ortega, PhD, Arturo Vargas Bustamante, PhD, and Dylan H. Roby, PhD

ABOUT THE AUTHORS

Alexander N. Ortega is with the Department of Health Management and Policy, Dornsife School of Public Health, Drexel University, Philadelphia, PA. Arturo Vargas Bustamante is with the Department of Health Policy and Management, Fielding School of Public Health, University of California, Los Angeles. Dylan H. Roby is with the Department of Health Policy and Management, School of Public Health, University of Maryland, College Park.

 See also Ro et al., p. 2019.

In 2017, approximately 10.5 million undocumented immigrants lived in the United States. Although the number of undocumented immigrants has declined over the past decade, this estimate is triple that of the population size in 1990, when there were 3.5 million undocumented immigrants.¹ According to a recent population-based study of California residents, undocumented immigrants exhibit worse patterns of health care access and use than US-born residents or immigrants who are naturalized or hold green cards.² This is not necessarily surprising given that undocumented immigrants have largely been left out of policies aimed at improving insurance coverage and access to care.

Some states, including California, have used state Medicaid programs or organized county indigent care programs to care for undocumented youths and young adults. However, most states have excluded undocumented

immigrants from public insurance programs, and they are explicitly excluded from the marketplace exchanges and the federally funded Medicaid expansions provided by the Patient Protection and Affordable Care Act.^{3,4}

In their article in this issue, Ro et al. (p. 2019) used administrative data from the Los Angeles County + University of Southern California Medical Center to describe differences in illness severity, length of hospital stays, and repeat hospitalizations between undocumented immigrant patients and full-scope Medi-Cal (California's Medicaid program) patients. To classify undocumented immigrants, they used as a proxy restricted-scope Medi-Cal, which is limited to emergency and pregnancy care for low-income Los Angeles residents who meet the Medi-Cal income threshold but do not meet immigration status requirements (e.g., US national, lawful permanent resident).

Their main findings were that younger (18–64 years) undocumented immigrant patients had less severe illness and spent less time in the hospital than younger Medi-Cal patients, whereas older (≥ 65 years) undocumented patients also had less severe illness but had lengths of stay that were similar to those of older Medi-Cal patients. These findings generally confirm previous reports indicating that undocumented immigrants are unlikely to burden the safety net because of their healthy profiles and underuse of health services.^{2,5}

This study is an important contribution to the growing literature on the health and health care needs of undocumented immigrants. It is also one of the first studies in which administrative data have been used to understand patterns of hospital care use and illness severity among undocumented immigrants requiring hospitalization. The study, however, must be put into the context of the broader empirical literature on the topic.

For example, the authors used data from patients needing hospitalization from a safety net hospital in Los Angeles that serves a very vulnerable population with complex health needs, including many low-income patients on Medicaid or without insurance and a large homeless population. Thus, the observations are not necessarily representative of the undocumented population in Los Angeles County, the state of California, or nationally. Also, the strategy to identify undocumented immigrant patients through restricted-scope Medi-Cal creates potential selection bias among patients who are poor and have high medical needs. The authors attempted to account for this problem by controlling for homeless status and using an inverse probability weighted regression adjustment;

however, the unobserved differences between the undocumented immigrant and Medi-Cal populations are still likely to misrepresent the health advantages of undocumented immigrants.

Public health researchers have long observed a health advantage for recent immigrants, an advantage commonly referred to as the “healthy immigrant effect.”⁶ This phenomenon is the notion that recent immigrants are in better health, on average, than US-born residents or immigrants who have been in the country for a long period of time. Because access to and use of health care tend to be poor among undocumented immigrants, it has been assumed that in general they are at high risk for poor health outcomes; however, others have posited that they are protected as a result of the healthy immigrant effect.⁷

Indeed, Ro et al. concluded that despite poor access to care, undocumented immigrants had less severe illness than their Medi-Cal counterparts. It should be noted, however, that their measure of health trends was based on illness severity as measured by the relative risk of mortality among hospitalized patients. Most researchers assessing the health advantages of immigrants have used measures such as self-rated health status, physician-diagnosed chronic diseases, and health behaviors.^{2,8,9}

In 2017, it was estimated that 66% of undocumented immigrant adults in the United States had been in the country for more than 10 years, as compared with 41% in 2007. With acculturation and more time spent in the United States, it has been observed that the advantages of the healthy immigrant effect decline.¹⁰ Undocumented immigrants, particularly those who have been in the country for many years, would in theory benefit less from the

healthy immigrant effect. A recent California study showed that the immigrant health advantage with respect to cardiovascular behavioral health risk did not apply to undocumented immigrant Latino men whose health behaviors were similar to those of US-born Latino men; however, health patterns were better among undocumented immigrant Latinas.⁸ Similarly, a study of Latinos in Los Angeles County revealed that undocumented immigrants who had been in the United States for short durations had worse self-reported health than the US-born individuals.⁹

In 2018, it was estimated that 880 000 undocumented immigrants lived in Los Angeles County and that 680 000 (78%) of them were Latino.¹¹ In a study of California Latino immigrants, only 25% of undocumented immigrants reported being in excellent or very good health, as compared with 49% of US-born Latinos and 36% of naturalized Latino immigrants, even though undocumented immigrants were less likely to report physical health problems such as obesity, high blood pressure, asthma, and diabetes.² The lower odds of undocumented immigrants ever having been told by a provider that they had a physical health condition were a function of their having significantly worse access to care than other groups.³ Furthermore, some of the discrepancy, at least with respect to self-rated health, might be attributed to the measure’s response categories and the Spanish translation of “fair.”¹²

Immigrants have long been underserved in the US health care system, particularly those who are uninsured or do not have legal authorization status. The recent expansion of Medicaid coverage to undocumented immigrants up to the age of 26 years in California and new indigent care programs such as

MyHealthLA are examples of initiatives at the state and local levels that can partly address health care inequities among immigrants. Policy efforts to expand health insurance coverage to the growing aging undocumented immigrant population should include assessments of the situation in California, where the state legislature recently decided to expand Medi-Cal coverage to low-income undocumented immigrants 50 years or older. Undocumented immigrants 26 to 49 years old without insurance will continue to have some level of access to primary, specialty, and inpatient care through California’s indigent care programs, although these programs do not offer the same financial protection as private insurance or Medi-Cal.

Public health research on the health and health care of undocumented immigrants in the United States is quickly evolving. The Ro et al. article augments what has been observed in population-based studies, and it raises some interesting areas of inquiry for the field. The extent to which the healthy immigrant effect applies to undocumented immigrants, especially as they live in the country for longer durations and a growing proportion of them are aging, is still open for debate and discovery.⁷ We also need a better understanding of the health care needs of undocumented immigrants to inform policies and programs aimed at addressing health inequities. As the demography of the immigrant population continues to shift, it is imperative that public health research, practice, and advocacy focus on ways to engage all immigrants, especially those who are undocumented. [AJPH](#)

CORRESPONDENCE

Correspondence should be sent to Alexander N. Ortega, PhD, Drexel University, Dornsife School of Public Health, Department of Health Management

and Policy, 3215 Market St, Nesbitt Hall, Philadelphia, PA 19104 (e-mail: ano37@drexel.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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Public Health Under Siege: Improving Policy in Turbulent Times

Edited by: Brian C. Castrucci, DrPH, Georges C. Benjamin, MD, Grace Guerrero Ramirez, MSPH, Grace Castillo, MPH

This new book focuses on the importance of health policy through a variety of perspectives, and addresses how policy benefits society, evidently through increased life expectancy and improved health. The book describes how detrimental social determinants can be to the overall population health and emphasizes how the nation is centered on policy change to create equal health care opportunities for all sectors of health.

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The High Cost and Unequal Cancer Burden of Poor Diet in the United States

Marjorie L. McCullough, ScD, RD, and Farhad Islami, MD, PhD

ABOUT THE AUTHORS

Marjorie L. McCullough is with the Department of Population Science, American Cancer Society, Kennesaw, GA. Farhad Islami is with the Department of Surveillance and Health Equity Science, American Cancer Society.

See also Wang et al., p. 2008.

Poor diet is known to increase cancer risk and mortality, both directly and mediated by excess body fatness.^{1,2} Characterized by higher intakes of red and processed meat and sugar sweetened beverages (SSBs) and a lower intake of whole grains, vegetables, and fruit, poor diet is associated with the risk of multiple cancers, especially colorectal cancer.^{1,2} The most recent estimates of the proportion of new cancer cases attributable to suboptimal diet range from 4.2%³ to 5.2%,⁴ equating to approximately 67 000 to 80 000 total cases per year in the United States. Disparities in cancer outcomes among racial and ethnic groups and by socioeconomic status are well documented.^{5,6} For example, colorectal cancer death rates among Black males is more than 44% higher than among White males.⁵ Few studies, however, have quantified the impact of poor diet on cancer outcomes in these subgroups.

In this issue of *AJPH*, Wang et al. (p. 2008) estimate the lifetime health and economic costs of cancer attributable to poor diet among US subpopulations defined by age, sex, race/ethnicity,

education, income, and Supplemental Nutrition Assistance Program (SNAP) participation. The authors projected both direct and indirect effects of poor diet, the latter derived from a pooled analysis of three prospective cohorts, to obtain diet–body mass index associations. Additionally, using publicly available data and modeling techniques, the authors estimated that the impact of suboptimal intakes of seven dietary factors (whole grains, dairy products, fruits, vegetables, red meat, processed meat, and SSBs) accounted for 3.04 million new cancer cases (7.4%) and 1.74 million cancer deaths (7.7%) among US adults over a lifetime. These percentages are higher than previous estimates, likely because of the authors' use of lifetime modeling, whereas estimates from previous studies were for a specific year (i.e., 2014³ or 2015⁴). The authors also estimated \$254 billion in medical costs attributable to poor diet (7.8% of direct medical costs of 15 diet-related cancers in the United States).

A valuable contribution by Wang et al. is the quantification of disparities in the diet-attributable burden of cancer in subpopulations by cancer type and by

dietary factor. For example, the authors estimated that non-Hispanic Black persons would experience more diet-attributed incident cancers and cancer deaths than would non-Hispanic White persons, with a difference of 110 cases and 214 deaths per 100 000. With the exception of diet-attributed oral, pharyngeal, and laryngeal cancers, which was highest among White persons, this disparity applied to all cancers studied, but particularly colorectal cancer. This higher burden was greatest for low consumption of dairy and whole grains and for excess consumption of processed meats and SSBs.

Having attained less than a college education was also associated with a higher number of diet-attributable cancer cases and deaths compared with adults with a college degree (excess cases and deaths of 180 and 132 per 100 000, respectively). The largest differences were seen for colorectal cancer and applied to all dietary factors, most notably SSBs. Low family income and SNAP participation were associated with higher diet-attributable cancer cases and deaths, with the largest disparities related to consumption of SSBs and low consumption of whole grains. Approximately 72% of the cases associated with poor diet were attributed to direct diet–cancer etiologic effects, with the remainder mediated by body mass index.

ORIGINS OF DISPARITIES

The reasons for disparities in diet-related health outcomes are multifactorial and complex, with major aspects rooted in structural and social marginalization and discrimination, which result in poorer living environments, limited educational and occupational opportunities, and lower incomes

among people of color and other historically marginalized populations.⁷ Consequently, these populations often experience higher levels of unhealthy food marketing and lower access to healthy foods because of a lack of financial and other resources, including transportation insecurity. This is compounded by limited access to supermarkets with healthy, affordable, high-quality choices but high availability of low-cost, poor quality “fast foods” and less healthy foods from convenience stores.¹ Even when supermarkets are available, healthier foods are often more expensive.⁸ Disconcerting trends of widening gaps in diet quality have been documented by education and income.⁹

PUBLIC HEALTH IMPLICATIONS

Eliminating health inequities through addressing structural racism and improving social determinants of health, which influence factors such as education, wages, housing security, and access to medical care, will need greater societal efforts to address these fundamental causes of poor health. However, some nutrition interventions may help reduce disparities in ensuring food and nutrition security so that individuals have a fair chance at a healthy diet pattern. Several policy implications and approaches are enumerated by Wang et al. For example, the authors suggest priority targets for behavior change and policy strategies to reduce these disparities, including incentives and disincentives to steer SNAP food purchases and improving access to whole grains, fruits, and vegetables, as well as disincentives, such as taxation or labeling, to discourage SSB and processed meat consumption. SSB

consumption emerged as a key target in this work; various policy approaches—including taxation—have proven successful in several countries, regions, and cities,¹⁰ although few places in the United States have made progress in adopting these policies in the past few years.

Improving availability of drinking water and affordable, healthy, culturally appropriate foods in low-income and racial/ethnic minority communities are other approaches to addressing these gaps. Produce prescription programs and financial incentives for purchasing fruits and vegetables among SNAP participants are showing promise in US-based studies.¹¹ In these programs, when a person purchases produce with SNAP funds, they receive additional funds to spend on fruits and vegetables. Medically tailored meals are emerging as a way to support risk reduction and disease self-management. Reducing the marketing of unhealthy foods and beverages has also shown promise as a strategy. Finally, to inform policy and program efforts, there is a need for more implementation research on interventions that aim to reduce disparities.

STUDY STRENGTHS AND WEAKNESSES

The study by Wang et al. provides important evidence on diet-related cancer and economic disparities that can be used to inform future efforts to improve health equity. Despite its many strengths, this study also has limitations, some of which are acknowledged by the authors. For example, they used data on diet in people aged 20 years and older from surveys conducted in 2015 to 2018 to estimate diet across the life course, not taking into account

any potential changes in diet over different periods or the effects of early life diet on cancer outcomes, although the latter has not been extensively investigated in other similar studies either. In addition, more research is needed on indirect effects of diet on cancer through body mass index and on any differences in associations between diet and cancer risk by race/ethnicity; in this study, the authors used the same estimates for all racial/ethnic groups.

STUDY TAKEAWAYS

Food insecurity, at high levels in the United States because of the COVID-19 pandemic, is likely to exacerbate already existing racial and economic health inequalities. If left unaddressed, the disparities in diet quality, the disparities in the proportion of diet-preventable cancers, and the high economic costs identified by Wang et al. will continue or worsen. Innovative and societal solutions are required at national, state, territorial, tribal, and local levels to make it possible for all individuals to have the ability to follow dietary recommendations and cancer-prevention guidelines for optimal health and quality of life.

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CORRESPONDENCE

Correspondence should be sent to Marjorie L. McCullough, Department of Population Science, American Cancer Society, 3380 Chastain Meadows Pkwy. NW, Suite 200, Kennesaw, GA 30144 (e-mail: marji.mccullough@cancer.org). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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CONTRIBUTORS

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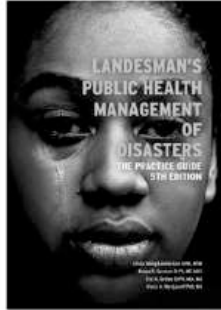
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Capitalizing on Missed Opportunities for Sexual Health Workforce Development by Adoption of a Sexual Health Paradigm

Vincent Guilamo-Ramos, RN, PhD, MPH, Adam Benzekri, MS, Marco Thimm-Kaiser, MPH, Amy Geller, MPH, Aimee Mead, MPH, Charlotte Gaydos, DrPH, MPH, Edward Hook, MD, and Cornelis Rietmeijer, MD

ABOUT THE AUTHORS

Vincent Guilamo-Ramos, Adam Benzekri, and Marco Thimm-Kaiser are with the Center for Latino Adolescent and Family Health, Duke University School of Nursing, Durham, NC. Amy Geller and Aimee Mead are with the Health and Medicine Division, National Academies of Sciences, Engineering, and Medicine, Washington, DC. Charlotte Gaydos is with the Division of Infectious Diseases, Johns Hopkins University School of Medicine, Baltimore, MD. Edward Hook is with the Department of Medicine, University of Alabama at Birmingham. Cornelis Rietmeijer is president of Rietmeijer Consulting, Denver, CO.

Sexually transmitted infections (STIs), a longstanding problem for Americans, were characterized as a hidden epidemic in a 1997 Institute of Medicine report.¹ More than two decades later, the National Academies of Sciences, Engineering, and Medicine (National Academies) have revisited the topic and released *Sexually Transmitted Infections: Adopting a Sexual Health Paradigm*,² a consensus study report with recommendations for responding to persistently high, increasing, and now record-level rates of reported STIs.³ Despite more than two decades of separation, the two reports discuss surprisingly similar key problems. STIs remain a major cause of morbidity, disproportionately affecting younger persons and having lifelong consequences.^{1,2} Reportable STI rates have

increased since 1997, and the latest data reflect an all-time high.³ Estimates suggest that approximately one in five people in the United States had an STI on any given day in 2018.⁴ Although the STI burden is increasing across all population groups, adolescents and young adults, women, men who have sex with men, and other groups underserved by mainstream health and public health systems remain disproportionately affected. Therefore, the dire need for increased public health attention and resources for addressing the “hidden epidemic” of STIs persists today.^{1,2}

The lack of progress in STI prevention and control is owing to longstanding underinvestment in the broader public health system and its workforce, as highlighted during the COVID-19

pandemic. For example, the Centers for Disease Control and Prevention’s (CDC’s) inflation-adjusted budget for STI prevention decreased by 40% since 2003.⁵ This underinvestment has contributed to the deterioration of the STI-focused public health infrastructure, programming, and workforce, including unmet staffing needs in state and local STI programs and declining availability of public health STI specialty clinics, an important safety net for individuals without health insurance.^{2,5}

We draw on the National Academies’ new report to outline a roadmap for rebuilding and expanding a workforce that is equipped for effective STI prevention and control, thereby addressing a missed opportunity from more than two decades ago: investing in STIs and sexual health as a public health priority.

ADOPTING A SEXUAL HEALTH PARADIGM

As a guiding paradigm for STI prevention and control, the National Academies committee that wrote the report recommends adopting a holistic perspective on sexual health as a key dimension of healthy living.² Wider recognition of this concept is reflected in calls from Congress to establish a Federal Office of Sexual and Reproductive Health and Wellbeing.⁶ Importantly, this new paradigm requires consideration of the sociostructural determinants of sexual health and STI risk beyond individual behavior. A holistic sexual health paradigm is supported by empirical evidence, which suggests that sociostructural factors are central in determining individual STI risk and in producing STI disparities, rendering health inequity a driving force in STI epidemiology.²

The proposed sexual health paradigm for STI prevention and control is more than a change in terminology and is a twofold opportunity for moving beyond enduring stigma and chronic underresourcing. First, adoption of a health-centered perspective moves toward the abandonment of the disease-focused, stigma-loaded framing of STIs that had persisted across previous nomenclature changes (i.e., “venereal disease” to “sexually transmitted disease” to “sexually transmitted infection”). Second, recognition of sexual health as an integral component of broader health and well-being creates opportunities for using additional resources and partnerships (e.g., in education, family services, community health) to supplement STI-specific funding and infrastructure and a STI-specific workforce. Notably, the COVID-19 response demonstrated the benefits of braiding and blending funds to facilitate cross-agency collaboration in improving health-related services.⁷

In relation to the workforce, the adoption of a sexual health paradigm has implications for (1) broadening ownership and accountability for STI prevention and control, (2) redefining STI specialists from a public health perspective, and (3) creating new sexual health-focused workforce opportunities across the public health system.

BROADENING OWNERSHIP AND ACCOUNTABILITY

There is a longstanding misperception by too many primary care practitioners that addressing sexual health and STIs is not their job, narrowly vesting responsibility for sexual health promotion and STI services in STI clinics or among a relatively small number of STI and sexual

and reproductive health specialists. However, additional practitioners across clinical health care and public health, most notably primary care providers, can be used if a well-being-focused sexual health paradigm is adopted and applied to workforce development. Therefore, the committee identifies a wide range of professionals and stakeholders as part of the sexual health workforce.

Although primary care providers (i.e., primary care physicians, nurse practitioners, and physician assistants), nurses, and many clinical behavioral health professionals are well positioned to deliver or facilitate sexual histories, STI vaccination, routine testing, and treatment, clinical health care generalists often do not sufficiently prioritize sexual health services.² Recognition of the important role of general practitioners in sexual health and STI services is required. The committee therefore recommends that clinical practice guidelines and training curricula for health care generalists define a minimum set of sexual health competencies, more heavily emphasizing the importance of the consistent delivery of recommended sexual health services, such as sexual histories, STI screening, and vaccination.²

As first-line providers trained to deliver most aspects of sexual health promotion, STI prevention, and STI management and as the largest segment of the health care workforce, nurses are particularly well positioned to increase the reach of sexual health services.² The committee encourages a broader scope of nursing practice in sexual health services as meaningful for strengthening the sexual health workforce and reducing STI disparities—guidance aligned with the vision outlined in the National Academy of

Medicine's *The Future of Nursing 2020–2030* report to use nurses for addressing social determinants of health and population health in the United States.⁸

Given that about 90% of the US population lives within two miles of a community pharmacy, pharmacists can serve as convenient entry points into the health care system, including for sexual health services.² The committee therefore highlights the utility of incorporating pharmacists into the sexual health workforce, particularly for STI testing using point-of-care tests.

Research to develop and improve biomedical tools, behavioral interventions, and sexual health service delivery models and implementation strategies to adopt existing evidence-based programs are essential for improving the response to STIs.² The report emphasizes the need for increased investments in sexual health research, including training programs and incentives for researchers to take up sexual health-focused work.

Research demonstrates that parents, health educators, community health workers, and civic and religious leaders can be influential in family- and community-based sexual health education and promotion.² It is essential to recognize these stakeholders as an important segment of the sexual health workforce and to mobilize the public health system to provide them with resources, training, and support.

Disease intervention specialists reduce the transmission of STIs by providing ground-level STI education and partner services, including counseling, testing, and referrals.² The committee reaffirms the importance of disease intervention specialists in promoting sexual health and addressing STIs, especially syphilis, and highlights their

utility for responding to outbreaks of other communicable diseases, for example through contact tracing for COVID-19.² The recently announced governmental investment of \$1.13 billion from the American Rescue Plan Act to strengthen the disease intervention specialist workforce to address COVID-19 and other infectious diseases is an important step in the right direction.⁹

REDEFINING SPECIALISTS

Beyond broadening ownership and accountability for sexual health, the development of sexual health and STI specialty expertise in all sectors of the public health system is integral. Traditionally, professionals considered as STI specialists have been mainly limited to clinical STI specialty providers and disease intervention specialists. Yet expertise across clinical, epidemiological, sociobehavioral, pharmaceutical, diagnostic, programmatic, and policy domains of public health is necessary to develop and implement a comprehensive strategy for sexual health promotion and STI control.²

For example, clinical sexual health specialty providers are needed to manage complicated STI cases and to provide technical assistance for health care generalists. Epidemiologists with sexual health specialty expertise are required to design and improve sexual health surveillance systems and to identify risk and protective factors to be used for behavioral, biomedical, diagnostic, and structural interventions. Practitioners and researchers with sexual health specialty expertise can implement existing and develop new interventions and advance the accuracy, timeliness, and practical utility of diagnostics. Public health administrators and public officials with sexual health

specialty expertise are crucial for designing and implementing public health programs and federal, state, and local regulations that facilitate, rather than obstruct, sexual health promotion and STI prevention and control. The group of professionals considered sexual health specialists should therefore be redefined by considering each relevant sector of the public health system.

CREATING OPPORTUNITIES

Given that professionals with sexual health specialty expertise remain scarce in many sectors of the public health system, efforts to attract and sustain such a workforce are needed, for example, through investments in sexual health–focused training opportunities and subsequent career paths. Various partners can be engaged to improve sexual health–focused training and career opportunities, including the CDC, state and local health departments, federally qualified health centers, the National Coalition of STD Directors, the National Network of STD Clinical Prevention Training Centers, disease intervention specialists training centers, universities (e.g., schools and departments of medicine, nursing, the biomedical sciences, public health, and social work), and commercial pharmaceutical businesses through public–private partnerships.^{2,10}

Both the existing workforce and the next generation of professionals require opportunities for the development of sexual health and STI specialty expertise. For this workforce, sexual health and STI-focused continuing education is scarce, and additional opportunities should be offered by the CDC, state and local health departments, STI and HIV expert providers, and STI

prevention training centers.² For future professionals, sexual health–focused electives offered as part of professional education for physicians, nurses, physician assistants, and other health and social service professionals (e.g., pharmacists, behavioral health practitioners, social workers, public health professionals) are important (e.g., as sexual health certificate programs as extensions to professional degrees).² In addition, the expansion of fellowship-based sexual health specialty training programs could strengthen the pipeline for future clinical, epidemiological, programmatic, and policy leaders.²

The committee emphasizes that the Department of Health and Human Services, especially through the CDC Division of STD Prevention, has particular responsibility for leadership in elevating STIs on the national public health agenda by promoting sexual health–focused training and career opportunities. To this end, the involvement of and support for local stakeholders in planning and coordinating sexual health promotion and the responses to STIs at the jurisdiction level is crucial. To develop effective local sexual health workforces around the country, the report envisions a formal infrastructure for training and technical assistance that draws on existing sexual health and STI expertise from STI and HIV expert providers; for example, through local or regional STI resource centers.

CONCLUSIONS

The 1997 *Hidden Epidemic* report was a missed opportunity to invest in STIs as a national public health priority. The new National Academies' report on STIs provides another opportunity to address this continuing challenge,

particularly given that the release of the first National STI Strategic Plan for the years 2021 through 2025 indicated renewed interest in STIs at the federal level.¹¹ The National Academies' report defines sexual health as an integral component of healthy living and as a guiding paradigm for future STI prevention and control efforts. This new sexual health paradigm necessitates broadening ownership and accountability for STI prevention and management to include practitioners and stakeholders not traditionally considered part of the sexual health workforce, redefining STI specialists from a public health perspective, and creating opportunities for sexual health specialty training and careers across the public health system.

Decades of underinvestment in the broader US public health system have contributed to record levels of reportable STIs. The COVID-19 pandemic has increased momentum for greater prioritization of the public health infrastructure and workforce.^{2,12} Too often, however, past calls to strengthen infrastructure and the workforce in response to contemporary public health priorities have yielded only inadequate and short-lived action, such as the response to the 1997 *Hidden Epidemic* report. Using the renewed focus on public health to address one of the leading cause of morbidity in the United States—namely STIs—through targeted workforce development is an opportunity that should not be wasted.

AJPH

CORRESPONDENCE

Correspondence should be sent to Vincent Guilamo-Ramos, Center for Latino Adolescent and Family Health, Duke University School of Nursing, DUMC 3322, 307 Trent Dr, Durham, NC 27710 (e-mail: vincent.ramos@duke.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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Caregiving Across Cultures: Priority Areas for Research, Policy, and Practice to Support Family Caregivers of Older Asian Immigrants

Minakshi Raj, PhD, MPH, Sasha Zhou, PhD, Stella S. Yi, PhD, and Simona Kwon, DrPH

ABOUT THE AUTHORS

Minakshi Raj is with the Department of Kinesiology and Community Health, University of Illinois at Urbana-Champaign, Champaign. Sasha Zhou is with the Department of Public Health, Wayne State University, Detroit, MI. Stella S. Yi and Simona Kwon are with the Department of Population Health, New York University, New York, NY.

About 4.6 million older adults (aged 60 years and older) in the United States are foreign born, and Asian Americans are projected to become the largest immigrant group in the United States by 2055.¹ Older Asian immigrants have to navigate new sociocultural contexts, including relationships with health care providers, dietary recommendations and adjustments, and care-seeking norms.² They may also experience structural challenges, such as discrimination in the health care system because of race, ancestry, and language.³ These adults often require assistance from unpaid family caregivers (e.g., adult children)—with whom they are also likely to reside—to navigate these sociocultural complexities and barriers in the health care system.⁴ Yet, our understanding of the experiences and challenges of family caregivers of older Asian immigrants remains limited despite their unique circumstances as children of

immigrants who are simultaneously balancing and navigating multiple cultural identities of their own while often being viewed as perpetual foreigners or outsiders to US culture.⁵

Immigration history and experiences likely also influence the use of formal services. Minority family caregivers are less likely than White caregivers to use formal support services (e.g., mental health treatment), suggesting that they may face additional burdens and barriers in the process of supporting their older relative in the health care system.⁶ Identifying ways to better support family caregivers as they care for their immigrant older relatives is critical for promoting inclusion in the health care system. This requires concerted research on Asian American caregivers and their experiences and challenges in their role supporting older relatives with household, daily functioning, and health care activities as well as navigating sociocultural aspects of care.⁷

It also requires recognition and consideration of the diversity of experiences and challenges faced by Asian American families, including heterogeneous Asian ethnic subgroups with distinct cultures, languages, needs, and preferences.

We have identified four priority areas to advance research, practice, and policy on the role of family caregivers of older Asian immigrants as informed by our ongoing qualitative research with 30 caregivers of older immigrants from Asian subgroups, including those identifying as South Asian, Chinese, Korean, and Vietnamese: (1) caregivers as sociocultural navigators in the health care system, (2) caregivers' role in detecting and protecting against discrimination, (3) addressing the mental health needs of caregivers and their relatives, and (4) recognizing the heterogeneity of experiences by cultural background.

SOCIOCULTURAL CONTEXT OF CARE DELIVERY

Family caregivers of older immigrants play a critical role in navigating the social and cultural context of health care delivery in the United States on behalf of, or with, their older relatives. This role may be mediated by the caregiver's own identity as an immigrant; for instance, caregivers who are themselves first- or 1.5-generation (i.e., immigrated in childhood) immigrants may have responsibilities with linguistic translation or interpretation of medical terminology, procedures, and paperwork in the absence of professional interpreters in the health care setting.⁸ However, second-generation (i.e., US-born) immigrants may struggle with

translating cultural phrases or idioms in the health care context. For example, one participant explained that the closest translation for herbal medicine in Hmong is the word for wood chips, which resulted in confusion during a health care visit. Caregivers also discussed their role in navigating the link between dietary practices in health care settings and culture and tradition (e.g., the impact of the Vietnam War on concerns about food security in long-term care). Caregivers may be tasked with interpreting their older relative's customs and practices to their relative's health care provider or helping their relatives reconcile US health care and health care-seeking norms (e.g., use of primary care services rather than emergency departments).

PROTECTING AGAINST DISCRIMINATION

In our study, family caregivers reported detecting discrimination or unfair practices experienced by their older relative, even if their relative normalized or did not recognize those incidents as being discriminatory. They are also attuned to the possibility of discrimination occurring at the intersection of race and ancestry, language proficiency, and gender. For instance, multiple caregivers recalled experiences in which their relative was ignored for hours despite going in with an urgent care need while other patients received care. In instances of observing discrimination, caregivers in our study felt responsible for finding their relative a new provider, and although only a few caregivers felt comfortable speaking up about an incident, many expressed subsequent discouragement and distrust of the health care system.

MENTAL HEALTH NEEDS

Family caregivers in our study discussed challenges with loneliness and isolation as they became increasingly responsible for serving the daily functioning, social, and mental health needs of their older relatives. Several did not recognize that they were a caregiver as a result of complex cultural circumstances, such as norms around filial piety, consequently limiting their use of support services.⁹ They also discussed observing and supporting their relative through mental health challenges that were undiagnosed because of a lack of culturally and linguistically appropriate screening tools. Asian Americans, for example, exhibit cultural variations in the way distress is expressed—called “idioms of distress” in the *Diagnostic and Statistical Manual of Mental Disorders, Fourth Edition*. These idioms of distress affect the likelihood and type of service utilization and the way mental health symptoms are reported in research. For example, an extensive line of research has found that East and Southeast Asians are more likely to experience mental health symptoms in somatic rather than psychological terms.¹⁰ However, somatic symptoms, which include headache and fatigue, are not included in most depression and anxiety screenings.

DISAGGREGATING DATA

Improving the quality of research and subsequent implications for health care and supports for family caregivers and their older Asian immigrant relatives requires disaggregation of data on Asian Americans from research conceptualization to analysis and interpretation.^{11–13} For example, many South Asian family caregivers of Indian origin

in our study discussed how they and their older relative follow a strict vegetarian diet, limiting dietary options in health care facilities. A Hmong caregiver explained that in Hmong culture meat is a symbol of wealth, and modifying her relative's diet at home to meet clinical recommendations to manage cardiometabolic health conflicts with this cultural norm. Policies and practices—such as dietary guidelines—that assume that needs and preferences of all Asian immigrants and their caregivers are uniform ultimately reinforce a lack of inclusion in the system.

RECOMMENDATIONS

Our ongoing work informs several recommendations. First, diversifying the health care workforce (e.g., social workers, community health workers, medical interpreters, and dietitians) and building cultural literacy across care teams is critical to supporting family caregivers in their responsibilities as a sociocultural navigator. Further, health systems should facilitate early relationships and rapport building between patients, family caregivers, and these health professionals to promote trust.

And yet, although representation in the workforce is a promising avenue for promoting cultural sensitivity, health care institutions must also recognize and address the fact that seeking care from the health care system does not necessarily mean that patients are truly included in the system.¹⁴ Maintaining these cultural barriers will only perpetuate a health system that lacks inclusion, thereby exacerbating health disparities. For instance, health care facilities could offer foods that cater to multicultural dietary preferences, and clinicians could recommend dietary modifications that are culturally

relevant. Consistently, there is a need to include multiple Asian American communities in research and to disaggregate their data to compare across communities. Further, racial and ethnic differences in outcomes should not be attributed merely to cultural identity but, again, to the sociocultural context in which older adults and family caregivers live as well.

Future research should also consider diverse methodologies (i.e., quantitative and qualitative), approaches (e.g., individual and dyadic interviews; longitudinal surveys), and tools (i.e., translation of surveys, interviews, and clinical trials—e.g., for dementia—in multiple languages). These considerations are crucial for enhancing the quality of research and the applicability of findings to practice and policy so that the distinct needs of different Asian American communities are considered rather than a having a “one size fits all” approach to policy. For instance, intake forms for adding prospective participants to registries should be available in multiple languages to ensure the availability of research opportunities for Asian immigrants and their family caregivers and the accessibility of study components. Scales used in clinical settings to assess psychological distress should be translated not only in a breadth of languages but also in a reliable and valid way that is considerate of the cultural context in which language operates.¹⁵ Support groups and interventions for older adults and family caregivers should recognize that cultural and associated social barriers may limit the inferences drawn from previous literature. These limitations should be acknowledged and considered when developing subsequent interventions.

CONCLUSIONS

Supporting Asian immigrants and their family caregivers is a health system imperative and requires further insight into the caregiving role, caregiver and immigrant mental health needs, cultural sensitivity and relevance of programs and research, and disaggregation of data to illuminate the varied experiences and associated challenges and needs of distinct Asian American communities. *AJPH*

CORRESPONDENCE

Correspondence should be sent to Minakshi Raj, 2007 Huff Hall, 1206 South Fourth St, Department of Kinesiology and Community Health, Champaign, IL 61820 (e-mail: mraj@illinois.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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CONTRIBUTORS

M. Raj conducted the research study. M. Raj and S. Zhou conceptualized the study. All authors wrote and reviewed the editorial.

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An Unnecessary Gift for COVID-19 Vaccines and Therapeutics: The Medical Countermeasure Priority Review Voucher

Reshma Ramachandran, MD, MPP, Ravi Gupta, MD, and Jing Luo, MD, MPH

ABOUT THE AUTHORS

Reshma Ramachandran is with the National Clinician Scholars Program, Yale School of Medicine, New Haven, CT, and the Veterans Affairs Connecticut Healthcare System, West Haven, CT. Ravi Gupta is with the National Clinician Scholars Program, University of Pennsylvania, Philadelphia, and the Corporal Michael J. Crescenz VA Medical Center, Philadelphia, PA. Jing Luo is with the Department of Medicine, University of Pittsburgh School of Medicine, Pittsburgh, PA.

When the US Food and Drug Administration (FDA) approved remdesivir in October 2020 for the treatment of COVID-19, it awarded the manufacturer, Gilead Sciences, an additional prize: a medical countermeasure priority review voucher. Upon full FDA approval, COVID-19 vaccine manufacturers may also be awarded this regulatory incentive. These vouchers are meant to motivate manufacturers to develop medical countermeasures (MCMs), which are defined as medical products that treat or prevent harm owing to a biological, chemical, radiological, or nuclear agent identified as a material threat, or a condition caused by administering a drug against such an agent.¹ However, given the significant public funding already invested in these products before approval for both their development and procurement, we believe these vouchers offer little additional incentive and are unnecessary. As manufacturers of COVID-19 vaccines and other

therapeutics seek full approval and become eligible for MCM priority review vouchers, their use should be reconsidered.

WHAT ARE THEY?

Priority review vouchers were originally introduced through the FDA Amendments Act of 2007 (Pub L No. 110-85) to incentivize pharmaceutical manufacturers to develop drugs or biologics for tropical diseases, which are largely concentrated in developing nations and offer limited profit potential, in exchange for expedited review of a future new drug application.² In 2012, priority review vouchers were extended to sponsors of FDA-approved drugs treating rare pediatric diseases. When a priority review voucher is redeemed by a manufacturer, it allows a product that would have otherwise received a “standard review” of 10 months to instead receive a “priority review” of six months, allowing earlier market entry.

This designation directs the FDA to ensure sufficient attention and resources to the product’s application so that the hastened regulatory review deadline of six months is met. Companies awarded this voucher can use it for a product in their portfolio or sell it to another company. Recent vouchers have sold for about \$80 to \$130 million.³

In 2016, the 21st Century Cures Act (Pub L No. 114-255) broadened priority review vouchers to sponsors of MCMs.¹ To be eligible for an MCM priority review voucher, the Department of Homeland Security (DHS) together with the Department of Health and Human Services (HHS) must issue a determination that a potential public health emergency, such as the COVID-19 pandemic, is a material threat.^{1,4} Although a list of material threats that informs the issuance of MCM priority review vouchers must be updated annually, no such list has been made publicly available since 2017.⁵ Remdesivir was the fourth medical product for which an MCM priority review voucher was awarded since the program’s inception in 2016 (Table 1). Although COVID-19 has not been publicly listed as a material threat, congressional committees and manufacturers have been notified of this designation, allowing manufacturers to apply for this additional incentive when seeking FDA approval for treatments and vaccines.

ARE THEY NECESSARY?

An examination of the existing limited literature by the Government Accountability Office reviewing each of the three available priority review vouchers (tropical diseases, rare pediatric diseases, and MCMs) generally did not find any effect of priority review vouchers on innovation, although several sponsors claimed the program affected

TABLE 1— Research and Development and Procurement Support for FDA-Approved Medical Countermeasures Having Received Priority Review Vouchers

Product and Commercial Sponsor	Material Threat	Discovery of Product (Sponsor)	Conducted Pivotal Trial (Sponsor)	Expedited Review Pathway	Exclusivity Granted (Years)	Public Procurement	Date MCM Priority Review Voucher Issued
TPOXX (tecovirimat)—SIGA Technologies, Inc.	Smallpox	NIAID (NIH) ⁶	SIGA (NIH and BARDA) ⁷	Fast-track, priority review	Orphan (7)	BARDA: Sep 2018: \$629.0 million ⁸	Jul 13, 2018
Jynneos (smallpox and monkeypox vaccine)—Bavarian Nordic A/S	Smallpox, monkeypox	Bavarian State Vaccination Institute (German Government) ⁹	USAMRIID (BARDA) ¹⁰	Priority review	Biologic (12)	BARDA: Sep 2017: \$299 for 13.0 million doses ¹¹ ; Apr 2020: \$202 million for 1.4 million doses	Sep 24, 2019
INMAZEB (atoltivimab, maffivimab, odesivimab-ebgn)—Regeneron Pharmaceuticals	Ebola	Regeneron (BARDA) ¹²	NIAID, WHO, and INRB in DRC, and African Coalition for Epidemic Research, Response, and Training (NIH) ¹³	Breakthrough therapy, priority review	Orphan (7); Biologic (12)	BARDA: Jul 2019: \$345 million for 6-y contract ¹⁴	Oct 14, 2020
VEKLURY (remdesivir)—Gilead Sciences, Inc.	COVID-19	NIAID (NIH) ¹⁵	NIAID (NIH) ¹⁶	Fast-track, priority review	New chemical entity (5)	HHS: Jun 2020: \$500 million for 500 000 doses ^{17,18}	Oct 22, 2020
STRATAGRAFT (allogeneic cultured keratinocytes and dermal fibroblasts in murine collagen-dsdt)—Stratatech Corporation	Thermal burns	University of Wisconsin-Madison (NIH) ¹⁹	StrataTech (BARDA) ²⁰	Regenerative medicine advanced therapy, priority review	Orphan (7); Biologic (12)	BARDA: Oct 2015: \$247 million for 5-y contract ²¹	Jun 15, 2021

Note: BARDA = Biomedical Advanced Research and Development Authority; DRC = Democratic Republic of Congo; FDA = US Food and Drug Administration; HHS = US Health and Human Services; INRB = Institut National de Recherche Biomédicale; MCM = medical countermeasures; NIAID = National Institute of Allergy and Infectious Diseases; NIH = National Institutes of Health; USAMRIID = US Army Medical Research Institute of Infectious Diseases; WHO = World Health Organization.

their development decisions.³ Moreover, one analysis found that 25 of the 26 MCMs in clinical trials received public funding for their development, prompting the study's authors to suggest that alternatives to the priority review voucher program might better motivate the development of these products.

It is difficult to determine whether a voucher is cost-effective, as development cost data are not publicly available. However, the fact that MCMs already receive significant public investment, regulatory incentives, and federal procurement contracts (Table 1) suggests that a voucher may be unnecessary. For current devastating threats like COVID-19 for which there is a high demand and therefore a viable market, granting a priority review voucher as an additional incentive may not be needed at all.

The original discovery of all five MCM products was publicly funded—the innovation of four products was underwritten by the US government and one by the German government. The US government also sponsored the late-stage clinical trials supporting FDA approval of all five products. For three products, US federal agencies designed and ran these trials.

Ahead of approval, the FDA also granted these MCM products additional regulatory incentives, including designations that allowed these drugs and vaccines to receive expedited review. Additionally, these MCM products received further intellectual property protections through issuance of exclusivity periods upon FDA approval, barring generic competition for 5 (for new chemical entities), 7 (for orphan drugs), or 12 (for newly approved biologics) years.

Besides funding their development, the federal government ensured manufacturers a market for these MCMs

through the purchase of large volumes of each product. Each advance purchasing agreement for these MCM products, valued at several hundreds of millions of dollars, was often secured before their approval. Effectively, the federal government paid twice for these products: first through direct funds for their development and then again for their procurement.

When applied, priority review vouchers also create an undue burden for patients by forcing the FDA to rapidly assess the safety and efficacy of future therapies that would not have received an expedited review designation otherwise. Moreover, an analysis of FDA approvals for drugs that have received such an expedited review designation found they involved fewer pivotal trials, enrolled fewer patients in their trials, and, more often, used surrogate endpoints in lieu of more clinically relevant ones.²¹ For example, in March 2018, Novartis redeemed a priority review voucher to expedite approval of siponimod for relapsing forms of multiple sclerosis.³ Although FDA reviewers questioned whether the pivotal clinical trial had “sufficient evidence to support the approval of siponimod with an indication for relapsing forms of MS”^(p13) and although the drug had a “relatively modest”^(p12) treatment effect, the drug was approved in less than six months.²²

POLICY RECOMMENDATIONS

Of course, addressing pandemics like COVID-19 requires rapid development of medical products. But at the onset of this pandemic, manufacturers quickly shifted focus to COVID-19 to develop novel drugs and vaccines, catalyzed by both immense research and financial

investments from the federal government and by extraordinary public health need, which ensured a viable market for these technologies. Additional regulatory incentives were created for these products, including expedited review through the Coronavirus Treatment Acceleration Program, a newly created special emergency FDA program.

Congress could allow the MCM priority review voucher program to expire on its scheduled date of October 1, 2023. However, ahead of this, the FDA will likely award full approval to COVID-19 vaccines and therapeutics eligible for an MCM priority review voucher despite having already received significant federal funding for the development, manufacturing, and procurement of these products.

Before the MCM priority review voucher program expiration, the DHS and the HHS should remove the “material threat” determination for COVID-19, preventing indicated vaccines and therapeutics such as those for COVID-19 from receiving the voucher upon approval. Pfizer and BioNTech, manufacturers of the only fully approved COVID-19 vaccine to date (Comirnaty),²³ did not receive an MCM priority review voucher, perhaps in acknowledgment that the voucher was unnecessary given significant public funding support for the vaccine’s development and procurement,^{24,25} as well as record profit due to high demand.²⁶ More broadly, Congress should amend the current statute to further narrow the eligibility criteria for medical countermeasures to ensure more judicious granting of the MCM priority review voucher. This could include the product having already received substantial public investment or other regulatory incentives for its

development or procurement or, if the product is indicated for an ongoing threat with high demand, making such a voucher unnecessary. Additionally, as with the five MCM products that received priority review vouchers, the federal government should instead ensure sustained funding for the development of other MCMs in exchange for adequate supply and reasonable pricing in lieu of vouchers.

Moving forward, Congress should reconsider the MCM priority review voucher program altogether, acknowledging that there is no evidence that it promotes the development of such products. For example, in the next iteration of the 21st Century Cures Act (Cures 2.0) or upcoming reauthorization of the Prescription Drug User Fee Act (Pub L No. 102-571) in 2022, legislators should sunset the MCM priority review voucher ahead of its expiration. Although intended to spur necessary innovation, there is no clear evidence of any added value from these vouchers in the face of substantial government investment for research, development, and procurement of treatments and vaccines to combat these serious threats. [AJPH](#)

CORRESPONDENCE

Correspondence should be sent to Reshma Ramachandran, Yale School of Medicine, 333 Cedar St, SHM IE-61, New Haven, CT 06510 (e-mail: reshma.ramachandran@yale.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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The Hemp Loophole: A Need to Clarify the Legality of Delta-8-THC and Other Hemp-Derived Tetrahydrocannabinol Compounds

Eric C. Leas, PhD, MPH

ABOUT THE AUTHOR

The author is with the Herbert Wertheim School of Public Health and Human Longevity Science, University of California, San Diego, La Jolla.

On June 22, 2021, Connecticut became the newest state to legalize recreational cannabis, raising the total number of US states allowing recreational cannabis to 18. One of the promissory notes of state-led legalization is that it will bring cannabis out of an illicit market and into a more transparent one with better safety standards. Meeting this challenge will require properly defining a group of compounds that fall into the category of tetrahydrocannabinol (THC) but can have different structures and origins. Here I discuss loopholes that allow for legal hemp production to yield a THC compound that has escaped state and federal regulation because of its hemp, rather than cannabis, origins. I make the case that the loopholes that allow THC compounds to be sold as hemp ought to be closed and provide a brief case study of one state that has done an exceptional job in closing such loopholes.

WHERE DO THC COMPOUNDS COME FROM?

The label THC refers to a group of compounds that in chemistry are called isomers. THC isomers have the same formula and structure but a different arrangement of atoms and potentially different pharmacological properties. The most abundant naturally occurring THC isomer is Δ^9 -tetrahydrocannabinol (delta-9-THC). Delta-9-THC is the main cannabis-specific compound (called a cannabinoid) responsible for mediating the psychotropic effects of cannabis. Because of delta-9-THC's ubiquity in most cannabis strains, it is often referred to universally as "THC," but other THC isomers can also naturally occur, just usually in smaller amounts.

One other THC isomer that naturally occurs in cannabis plants is Δ^8 -tetrahydrocannabinol (delta-8-THC). Delta-8-THC is nearly identical in chemical structure to delta-9-THC, differing

only by the location of a carbon-carbon double bond.¹ Because the structure of delta-8-THC is slightly different from that of delta-9-THC, it interacts with the human body in ways that can differ from delta-9-THC. For example, at the same dose, delta-8-THC produces a psychotropic effect that is similar to but slightly reduced from delta-9-THC. The pharmacological profile of delta-8-THC also suggests it has antiemetic, anxiolytic, appetite-stimulating, analgesic, and neuroprotective properties, indicating that it may have therapeutic applications and that some of these applications may differ from delta-9-THC.

Both delta-8-THC and delta-9-THC naturally occur in cannabis, but they can also be synthesized. One means of synthesis is conversion from another group of naturally occurring cannabis isomers called cannabidiol (CBD). Because CBD isomers are similar in structure to THC isomers, they can be converted to THC isomers through a relatively simple series of chemical reactions. The main method of converting CBD to delta-8-THC yields a solution containing delta-8-THC and delta-9-THC as well as other byproducts from the associated reactions. This solution can be further processed to remove delta-9-THC and then added to various consumer goods for consumption or application.

LOOPHOLES ALLOW SOME THC COMPOUNDS TO BE SOLD NATIONWIDE

Many manufacturers in the United States have made the argument that delta-8-THC products are "hemp products" and not "cannabis products," which, if true, affords them legal rights and privileges. Hemp is a variety of the

cannabis plant species in which 0.3% or less of its mass contains delta-9-THC when measured after it is dried. In the past three years, hemp and derivatives extracted from hemp have become widely available in the United States after a US congressional law called the 2018 Farm Bill removed hemp from the definition of marijuana in the Controlled Substances Act.² This has allowed hemp to be legally cultivated for its commercial and industrial purposes (e.g., to make paper, rope, and clothing) while retaining the schedule I classification of cannabis strains containing concentrations of delta-9-THC above 0.3%.

Because hemp contains very little delta-9-THC, it is not generally considered psychotropic; however, because it contains an abundance of CBD, it is possible to synthesize THC isomers from hemp. The processes used to convert CBD or other cannabinoids found in hemp to THC isomers were not explicitly prohibited in the 2018 Farm Bill or in most state laws governing hemp. One interpretation of the omissions of a broader definition of THC and the processes of synthesizing THC isomers from hemp laws is that this implies that THC isomers can be produced and sold legally under hemp laws so long as hemp and hemp derivatives are legally obtained and the final products contain 0.3% or less delta-9-THC.

These “hemp loopholes” have created a new marketplace for delta-8-THC products that uses sophisticated sourcing and distribution strategies designed to evade cannabis and hemp laws and appeal to consumers but also resemble a legitimate business. For example, one of the largest delta-8-THC manufacturers is based in Indiana, where both medicinal and recreational

cannabis products containing delta-9-THC remain prohibited.³ The manufacturer claims to obtain hemp and hemp extracts that are legally sourced from California, Colorado, and Oregon and then convert the extracts to other cannabinoids including delta-8-THC. The manufacturer sells bulk quantities of delta-8-THC solutions and an array of premixed delta-8-THC products that resemble traditional cannabis products (e.g., brownies, cookies, dabbing concentrates, gummies, vape cartridges) through its Web site and claims to have had these products featured in major news outlets such as ESPN, ABC News, and *Rolling Stone*.

The manufacturer also has a program to ship its products wholesale to retailers that resell them, as well as individually to consumers across the United States. To help guide purchasing decisions, the manufacturer provides a tool on its Web site to educate retailers and consumers on how it interprets federal cannabis and hemp laws and comparable laws in the 50 US states (Table 1). Nowhere on this manufacturer’s Web site does it suggest that it restricts shipments to specific places; on its “shipping” page, however, the manufacturer claims that “if, after ordering, we discover an item is illegal in your state we reserve the right to cancel and refund the order.”

CANNABIS CONTROL SYSTEMS PROMISE A SAFER MARKETPLACE

Hemp loopholes have created a marketplace in which delta-8-THC products are being widely sold outside of most of the cannabis control systems developed by states that have legalized recreational cannabis products. Although the marketplaces regulated by state-

run cannabis control systems remain unambiguously illegal at the federal level and have faced many challenges from the still-thriving illicit cannabis market, in theory they are designed to bring cannabis out of an illicit marketplace and into ones with more effective means for oversight and better safeguards that ensure consumer safety. Examples of some of the requirements in these systems provide a perspective on how they are intended to protect the public’s health in ways that the current marketplace for delta-8-THC does not.

First, most cannabis control systems have established a minimum purchasing age and require age verification to purchase cannabis products. Age requirements are designed to limit youth access to cannabis given that developing brains are more susceptible to addiction, exposure among youths potentially results in a greater number of heavy users, and heavy cannabis use among young people may have distinct health effects (e.g., cognitive impairments).⁴ For now, delta-8-THC retailers are setting their own minimum purchasing age (if any) and creating their own method of verifying the ages of their consumers.

Second, most cannabis control systems require cannabis products to be submitted to a certifying agency that tests them for potency, consistency, and a wide range of potential contaminants that are known to be used in illicit cannabis cultivation and processing. Without such requirements, adulterated products can enter the supply chain and harm consumers. For example, illicit cannabis vaping products containing the cutting agent vitamin E acetate were determined to be the primary cause of the 2019 outbreak of lung injury known as EVALI that led to

TABLE 1— Recreational Cannabis Laws and the Legal Status of Delta-8-THC Claimed by One Delta-8-THC Manufacturer as of July 22, 2021: United States

State	Recreational Cannabis Permitted	Delta-8-THC Permitted ^a
Alabama		Yes
Alaska	Yes	
Arizona	Yes	
Arkansas		
California	Yes	Yes
Colorado	Yes	
Connecticut	Yes	Yes
Delaware		
Florida		Yes
Georgia		Yes
Hawaii		Yes
Idaho		
Illinois	Yes	Yes
Indiana		Yes
Iowa		
Kansas		Yes
Kentucky		Yes
Louisiana		Yes
Maine	Yes	Yes
Maryland		Yes
Massachusetts	Yes	Yes
Michigan	Yes	Yes
Minnesota		Yes
Mississippi		Yes
Missouri		Yes
Montana	Yes	
Nebraska		Yes
Nevada	Yes	Yes
New Hampshire		Yes
New Jersey	Yes	Yes
New Mexico	Yes	Yes
New York	Yes	Yes
North Carolina		Yes
North Dakota		
Ohio		Yes
Oklahoma		Yes
Oregon	Yes	Yes
Pennsylvania		Yes
Rhode Island		
South Carolina		Yes
South Dakota		Yes
Tennessee		Yes

(continued)

2668 hospitalizations and 68 deaths.⁵ Delta-8-THC products have not been linked to EVALI, but such events expose the potential dangers of certain manufacturing processes that, if left unchecked, could harm consumers. Although some delta-8-THC manufacturers claim to test their products, the methods used to convert CBD to delta-8-THC have not been well studied with respect to quality assurance, and there are currently no established standards for testing delta-8-THC products for potentially toxic or otherwise harmful substances that could be byproducts of these processes.

Finally, most cannabis control systems have established detailed packaging and labeling standards including requirements for legible supplement information, warnings regarding the presence of psychotropic compounds, and restrictions on design features that could be attractive to youths (e.g., cartoons and names such as “kandyz”). Delta-8-THC products do not have such packaging and labeling requirements, and this could increase the risk of unintended use or overdose. The West Virginia Poison Center has already issued a warning after adults mistakenly consumed delta-8-THC products believing them to be CBD products and youths were hospitalized after consuming delta-8-THC gummies believing them to be candy.⁶

CLOSING HEMP LOOPHOLES

Precedent exists to incorporate delta-8-THC into federal regulations governing cannabis. According to the 1986 amendment to the Controlled Substances Act,⁷ a controlled substance analogue is a substance where “the chemical structure of which is substantially

TABLE 1— Continued

State	Recreational Cannabis Permitted	Delta-8-THC Permitted ^a
Texas		Yes
Utah		
Vermont	Yes	
Virginia	Yes	Yes
Washington	Yes	Yes
West Virginia		Yes
Wisconsin		Yes
Wyoming		Yes

Note. Blank spaces indicate that recreational cannabis or delta-8-THC is not permitted.

^aLegal status of delta-8-THC claimed by one delta-8-THC manufacturer.³

similar to the chemical structure of a controlled substance in schedule I or II.” The act also states that “a controlled substance analogue shall, to the extent intended for human consumption, be treated, for the purposes of any Federal law, as a controlled substance in schedule I.” Given that the chemical structure of delta-8-THC is nearly identical to that of delta-9-THC, it could certainly be considered a controlled substance analogue of delta-9-THC and enforced as such. The Drug Enforcement Agency has also promulgated that the use of any process that creates delta-9-THC as a byproduct at any point—as is the case during the conversion of CBD to delta-8-THC—is in violation of federal law.⁸ Although these rulings have been made, it is unclear whether the Drug Enforcement Agency has acted or will act on them.

In the current political landscape, federal enforcement of cannabis laws rarely occurs and will likely be inadequate to have any meaningful public health impact. Thus, state-by-state amendments to cannabis control laws are probably the best course of action to close hemp loopholes in a meaningful manner. Few states have anticipated

the potential for psychotropic compounds to be produced under hemp laws, but some have already taken actions to close their hemp loopholes in ways that put the public’s health first and serve as an example for other states.

The actions taken by Colorado represent an excellent model for closing a hemp loophole. First, Colorado’s controlled substance laws⁹ provide a clear definition of THC that broadly applies to isomers other than delta-9-THC, eliminating any ambiguity about whether delta-8-THC is considered an analogue to delta-9-THC. Second, the Colorado Department of Public Health¹⁰ and the Marijuana Enforcement Division¹¹ have made it clear to their stakeholders that the state considers the process of chemically modifying or converting naturally occurring cannabinoids from industrial hemp as noncompliant with the statutory definition of an “industrial hemp product,” clarifying that industrial hemp and hemp derivatives should not be used as precursors.

Third, the Marijuana Enforcement Division has also clarified the specific solvent and extract methods it allows and specified that the current

methods of converting cannabinoids do not fall in this category.¹¹ Finally, the Colorado Department of Public Health has explained that it is disallowing THC isomers produced by methods of converting cannabidiol or other cannabinoids from being added to food, dietary supplements, or cosmetics because there is insufficient evidence to determine whether any toxic or otherwise harmful substances are produced during the process of creating or converting THC isomers and may remain in the products ingested, applied, or used by consumers.

At present, these actions essentially ban manufacturers from producing or selling delta-8-THC or other THC isomers if they are converted from other cannabinoids in hemp or cannabis. However, Colorado’s Marijuana Enforcement Division has left open the possibility for stakeholders to work with regulators to examine potentially compliant methods to produce THC’s by converting cannabis-derived cannabinoids.¹¹ Colorado’s policies also do not appear to prohibit strains of cannabis that contain THC’s other than delta-9-THC or any compliant methods of extracting those compounds. This suggests that compliant delta-8-THC products could be sold within Colorado’s cannabis control system and that methods of converting cannabinoids could be used once their safety can be assured.

The actions taken by Colorado are commendable because they prioritize the public’s health while also acknowledging a potentially legitimate consumer demand for delta-8-THC and the potential value of safe methods of converting naturally occurring cannabinoids. For this reason, I believe that more states should follow in Colorado’s

footsteps and close their own hemp loopholes. *AJPH*

CORRESPONDENCE

Correspondence should be sent to Eric C. Leas, PhD, MPH, University of California, San Diego, 9500 Gilman Dr, Mail Code 0725, La Jolla, CA 94304-1334 (e-mail: ecleas@ucsd.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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Healthy Aging Through The Social Determinants of Health

Edited by Elaine T. Jurkowski, PhD, MSW
and M. Aaron Guest, PhD, MPH, MSW

This new book examines the link between social determinants of health and the process of healthy aging. It provides public health practitioners and others interacting with the older population with best practices to encourage healthy aging and enhance the lives of people growing older.

Healthy Aging: Through The Social Determinants of Health gives insight into the role each of these plays in the healthy aging process: health and health care; neighborhood and built environment; social support; education; and economics and policy.

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Accounting for the Full Scope of Health Consequences of Policy Decisions Across the Life Course

Sandro Galea, MD, DrPH, and Roger Vaughan, DrPH, MS

ABOUT THE AUTHORS

Sandro Galea is with the School of Public Health, Boston University, Boston, MA. Roger Vaughan is an AJPH associate editor and is with the Department of Biostatistics, The Rockefeller University, New York, NY.

🔗 See also Wang et al., p. 2008, Brenner, p. 1950, and Shen et al., p. 2027.

Our health is of immediate concern to all of us. Indeed, it is difficult to think of much that is of more immediate concern. Our health makes it possible for us to make choices about how we wish to live our lives. Clearly this statement is complicated by the many and varied expressions of health, and we recognize that many of us can live rich, fulfilled lives, even if in imperfect health. But fundamentally good health affords us the means to live however we wish, serving as a key element of our individual dreams and aspirations.¹

It is, then, not at all surprising that we tend to focus on health in the present. If we are unhealthy in the moment, it inhibits our capacity to do what we wish in the moment and compels us to seek a cure. This is when there is a synergy between our individual perception of health and the aspirations of medicine. Medicine is fundamentally concerned with restoring us to health, overcoming sickness in the short term so that we can return to living our lives as we choose to. That is a powerful and

compelling aspiration and explains, in no small part, why we as a country spend more on health care, by far, than any other comparable high-income country.² A health concern is an urgent concern, and we, as a country, have long judged it worth it to spend as much money as we can to ameliorate that concern, resulting in extraordinary spending on our health care.

And yet, as readers of this series of commentaries well realize, despite our spending on health care, our health as a country lags substantially behind our peer nations. Central to the reason for this mismatch between health spending and health outcomes is our underspending on preventing illness and, commensurately, on lessening the structural forces that contribute to illness throughout the life course.³ We underspend on lessening these structural forces for many reasons, but perhaps first among those is that it is often difficult to see the direct link between those forces and our health. Although we may be able to recognize logically that, for example, spending on parks

and recreation facilities will lead to opportunities for exercise and healthier populations in decades to come, that concern is readily drowned out by concerns with individual cardiac disease in the present, even if that cardiac disease could have been prevented with investment in structural forces in previous decades.

This, then, makes work that highlights the full scope of the life course a consequence of societal decisions that much more important. By documenting the health impacts of decisions in the present throughout our lives, such work builds on the scholarship of life course thinking^{4,5} and bears witness to the forces that generate the sum of health over many decades and may help us focus on structural causes of health.

HEALTH CONSEQUENCES

Three articles in this issue of *AJPH* do just that. The first article, by Wang et al. (p. 2008), documents the long-term health burden of suboptimal diets, focusing on cancer. Starting in 2017, they studied a closed cohort of US adults aged 20 years or older, modeling events across the lifespan. They found that suboptimal diets contribute to 3.0 million new cancer cases, 1.7 million cancer deaths, and more than \$250 billion in economic costs. The sheer scope of the consequences of suboptimal diets this study suggests is staggering, even more so when noting that many of the cancers studied were attributable to high consumption of sugar-sweetened beverages and low consumption of whole grains. Importantly, both sugar-sweetened beverages and diet composition are linked, relatively directly, to policy decisions that we make as a society.⁶ This means, in essence, that we are choosing foods available to us and thus

choosing to incur the long-term burden of cancer over the life course. When stated in such stark terms, these choices may seem inexplicable but perhaps are less so in the broader understanding of our focus on the immediate in health and our challenge, as discussed here, with thinking of the long-term consequences of actions today for our health tomorrow.

Our challenge with thinking about, and working to prevent, untoward health consequences in future decades is illustrated well in the current COVID-19 moment and is shown in two other articles. COVID-19 gripped the country and the world starting in 2019. Efforts to mitigate the spread of COVID-19 have included a wholesale limitation of population mobility, limiting employment for many, and limiting access to routine health care services for many more. An analysis by Brenner (p. 1950) shows how the economic consequences brought about by mobility restrictions related to slowing the spread of COVID-19 in the United States are associated with nearly 200 000 combined deaths in the long term as a result of unemployment and bankruptcies.

An article by Shen et al. (p. 2027) considers the interruption to regular health care services brought about by COVID-19. Using data from the Michigan immunization registry, this article shows overall declines in non-COVID-19 vaccination across the life course, including both pediatric and adult vaccinations. It will of course require substantial effort to catch up on vaccinations for all age groups to mitigate the health harms that can come from persons not being vaccinated. The potential long-term health challenges that have emerged from our dealing with COVID-19 do not suggest, of

course, that we should not have implemented the efforts needed to limit the spread of a novel, highly contagious pathogen, but they reinforce, with forceful clarity, the costs of such inaction over the life course. That should give us both pause and perspective and should inspire us to balance our actions in the immediate present with an awareness of their health costs in the coming decades.

ATTENDING TO FUTURE HEALTH

It is our job to focus on how we may improve population health over the life course. This requires an unyielding focus on what is difficult to keep in mind—the future—and a near constant reminder to those who are in a position to make policy decisions that will influence the long-term health consequences of our actions today. Articles such as the ones in this issue of *AJPH* do an invaluable job of building the evidence base that can help to this end. *AJPH*

CORRESPONDENCE

Correspondence should be sent to Roger Vaughan, The Rockefeller University, 1230 York Ave, New York, NY 10065 (e-mail: roger.vaughan@rockefeller.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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Moving Life Course Theory Into Action: Making Change Happen
EDITED BY SARAH VERBIEST, DRPH, MSW, MPH



Over the past decade, practitioners in the field of maternal and child health have gained a general understanding of Life Course Theory and its potential application to practice. This book focuses on moving Life Course Theory into practice, thereby filling a need for practitioners across a variety of fields and providing them with valuable strategies on how to apply this approach.

Moving Life Course Theory into Action is designed to fit into the busy lives of practitioners. With new ideas and strategies delivered in a compact handbook style format, each chapter includes key points that offer a quick summary of the main lessons advanced by the authors.

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Establishment and Evaluation of a Large Contact-Tracing and Case Investigation Virtual Training Academy

Olivia Golston, MPH, Michael Prelip, DPA, MPH, Debbie Bain Brickley, DrPH, MPH, Anne Cass, MPH, Lisa Chen, MD, Alina Dorian, PhD, Alice Gandelman, MPH, Chris Keh, MD, Andrew Maher, MPH, Roger Myrick, PhD, MA, Michael J. A. Reid, MD, MA, MPH, Karen White, MBA, MPH, Rachel Willard-Grace, MPH, and Shira Shafir, PhD, MPH

During the COVID-19 pandemic, the Virtual Training Academy (VTA) was established to rapidly develop a contact-tracing workforce for California. Through June 2021, more than 10 000 trainees enrolled in a contact-tracing or case investigation course at the VTA. To evaluate program effectiveness, we analyzed trainee pre- and postassessment results using the Wilcoxon signed-rank test. There was a statistically significant ($P < .001$) improvement in knowledge and self-perceived skills after course completion, indicating success in training a competent contact-tracing workforce. (*Am J Public Health*. 2021;111(11):1934–1938. <https://doi.org/10.2105/AJPH.2021.306468>)

Early in the COVID-19 pandemic, contact tracing and case investigation were identified as key strategies to help decrease the spread of disease. To support this public health need, it was necessary to rapidly build a workforce of contact tracers (CTs) and case investigators (CIs).

INTERVENTION

The goal of the Virtual Training Academy (VTA) is to train a skilled workforce of CTs and CIs to effectively combat the spread of COVID-19 in the state of California.

PLACE AND TIME

The VTA offered its first course in May 2020 and has held 35 weeks of introductory CT and CI training as of June 2021. Because the VTA is a virtual

training program, trainees and staff participate remotely.

PERSON

VTA trainees are prospective CTs and CIs. Although some are public health professionals, many are local or state government staff who have been redirected to support California's contact-tracing efforts during the pandemic. Most trainees have a non-health career background (Table A, available as a supplement to the online version of this article at <http://www.ajph.org>). VTA participants came from 56 of California's 61 local health jurisdictions and more than 100 different state departments.

PURPOSE

After he issued the March 19, 2020, stay-at-home order in California,¹ Governor

Gavin Newsom required statewide expansion of contact-tracing capacity to proceed with staged reopening.² The workforce is composed of two complementary roles. CIs contact individuals who have tested positive for COVID-19, provide instructions and resources for isolation, and elicit a list of contacts who may have been exposed to the virus.³ CTs then reach out to the identified contacts and instruct them to quarantine to prevent further disease transmission.³ With a population of nearly 40 million,⁴ California requires a large workforce of CTs and CIs to help contain outbreaks early at a local level. This need for rapid workforce development led to the establishment of the VTA.

IMPLEMENTATION

The VTA is implemented by subject matter experts from the University of

California, San Francisco and the University of California, Los Angeles in partnership with the California Department of Public Health. The flagship program of the VTA is a week-long introductory CT and CI training course, which the California Department of Public Health promotes to California’s local health jurisdictions. Originally, trainees could enroll in either a 14-hour CT track or an 18-hour CI track, depending on their anticipated job assignment. Both tracks covered the same core information,

with the CI track including additional information critical for the case investigation role. Given the demand for CIs throughout the state, the CT-only course was discontinued in December 2020.

During the course, trainees attend live didactic webinars covering topics that include the epidemiology of COVID-19, interviewing techniques, determining quarantine and isolation dates, and socioeconomic and racial disparities in COVID-19 cases and

deaths (Figure 1).⁵ Case reporting is covered briefly, although more in-depth training is provided outside the VTA through other trainings.

The VTA places a strong emphasis on active learning and skill building. Thus, to complement the webinars, all trainees participate in three hours (CT course) or six hours (CI course) of “skills lab” sessions. The skills labs are held in virtual breakout rooms and allow trainees to role-play interviews and discuss challenging scenarios with a small

Topic	CT Course	CI Course	Corresponding Knowledge Questions	Corresponding Self-Assessment Questions
Epidemiology of COVID-19	☑	☑	K1, K2	S1, S2, S3
Overview of CT/CI roles, responsibilities, and confidentiality	☑	☑	K3, K4, K5	S4, S5, S6
Principles of Containment	☑	☑	K7, K8	
Determining quarantine dates	☑	☑	K6	S7
Determining isolation dates		☑	K17	S12
Conducting a CT Interview	☑	☑	K9, K10	S8
Interviewing Skills & Building Rapport	☑	☑	K11, K12	S9
Conducting a CI Interview		☑	K16	S13, S14
Health Coaching Techniques	☑	☑	K13, K14	S11 (CT) or S15 (CI)
Cultural Humility	☑	☑	K15	S10

FIGURE 1— Overview of Virtual Training Academy Contact Tracing and Case Investigation Course Topics

Note. CI = case investigator; CT = contact tracer. See Tables B and D (available as supplements to the online version of this article at <http://www.ajph.org>) for specific knowledge check and self-assessment items.

group of peers under the guidance of trained course facilitators. The sessions are designed to develop skills and to raise trainee confidence. Optional skills labs are also offered, including a session to practice conducting interviews in Spanish.

To supplement synchronous course components, the course Web page offers resources to support trainees. Webinar recordings, presentation slides, and handouts are posted for trainees to review. Additionally, demonstration videos for the state's main case investigation and contact-tracing data management platform are provided.

EVALUATION

It is vital to California's COVID-19 response that the VTA successfully prepare trainees for the job. To evaluate this, pre- and postcourse assessments are administered.

The "knowledge check" contains true/false and multiple-choice questions that measure understanding of key course concepts (Figure 1; Table B, available as a supplement to the online version of this article at <http://www.ajph.org>). Those passing the postcourse knowledge check within two tries are eligible to begin work as a CT or CI.

The skills "self-assessment" measures self-perceived ability to perform key job skills (Figure 1; Table D, available as a supplement to the online version of this article at <http://www.ajph.org>). Self-perceived ability to perform each skill is rated on a scale from 1 (not at all) to 5 (very well). The ratings on all items are averaged to obtain a composite score of self-perceived skills for each trainee before and after the course.

We analyzed assessment results from the weeks of June 29, 2020 to March 22, 2021, during which time

1445 unique trainees completed the CT course and 3352 unique trainees completed the CI course. We did not include previous cohorts, as assessment instruments were not yet finalized. We included only trainees who completed all questions on both the pre- and postassessment in the analysis for that assessment so that the individual-level change in performance could be measured. We used a one-tailed Wilcoxon signed-rank test and found that the improvement in performance on the knowledge check was statistically significant ($P < .001$) for the CT and CI tracks, as well as for the subset of trainees with non-health career backgrounds in each track. A majority of trainees (76.9% in the CT track and 72.6% in the CI track) received a perfect score or missed only one question on the postcourse knowledge check, compared with 25.8% and 18.5% on the respective preassessments. The improvement in performance on the self-assessment was also statistically significant ($P < .001$) for both tracks and for the trainees who reported non-health career backgrounds. Overall results are shown in Table 1, and question-level results are shown in Tables C and D (available as supplements to the online version of this article at <http://www.ajph.org>).

These results indicate that VTA trainees left the VTA with a strong foundation in the knowledge required to be a CT or CI. Furthermore, they greatly improved their self-perceived ability to perform key job skills. The significant improvements extended to trainees without a health-related career background, who had lower mean pre-course scores than the other trainees but showed a larger percentage improvement. These results point to the effectiveness of the VTA at training

participants, regardless of career background.

ADVERSE EFFECTS

No adverse effects of the training program are known.

SUSTAINABILITY

Once an online training platform is established, courses can be deployed to fill gaps in the public health workforce and respond to emerging challenges. At the VTA, course offerings have evolved in response to the changing needs of health departments, with newer programs including specialized training courses for staff who are designated to assist schools with reopening or to manage outbreaks in congregate settings. The VTA has also recently launched a vaccine communication course intended to provide CTs and CIs with the skills necessary to address vaccine hesitancy. Although the frequency of program delivery may decline after the pandemic eases, virtual programs such as the VTA can be a centralized and sustainable way to train an effective public health workforce.

PUBLIC HEALTH SIGNIFICANCE

Based on these results, the VTA has succeeded in its goal of quickly and effectively training CTs and CIs who have a strong foundation in the knowledge and skills required for the job and will help to protect public health by reducing the spread of COVID-19.

Even once in-person trainings can safely resume, there are benefits to virtual training programs, including the ability to simultaneously train workforces in multiple geographic locations

TABLE 1— Summary of Assessment Results by Course Track and Career Background

	No.	Mean (Pre)	Mean (Post)	% Change	P
Knowledge check					
Contact tracing					
Overall	1404	12.17	14.03	15.4	< .001
Nonhealth	1054	12.03	13.98	16.2	< .001
Health	336	12.58	14.18	12.7	< .001
Case investigation					
Overall	3262	13.90	15.92	14.6	< .001
Nonhealth	1930	13.70	15.86	15.7	< .001
Health	1230	14.20	16.04	12.9	< .001
Self-assessment					
Contact tracing					
Overall	1327	3.19	4.33	35.6	< .001
Nonhealth	991	3.10	4.25	37.3	< .001
Health	321	3.48	4.56	31.2	< .001
Case investigation					
Overall	3079	3.33	4.45	33.9	< .001
Nonhealth	1827	3.24	4.40	35.9	< .001
Health	1157	3.46	4.54	31.3	< .001

Note. Analyses include trainees who answered all questions on the pre- and postassessment between June 29, 2020 and March 22, 2021. Knowledge check scores are out of 15 for the contact-tracing (CT) course and are out of 17 for the case investigation (CI) course. The self-assessment score is out of 5 for both tracks and is the average of 11 (CT) or 14 (CI) questions. See Tables B and D (available as supplements to the online version of this article at <http://www.ajph.org>) for specific questions. P values were obtained using a 1-tailed Wilcoxon signed-rank test comparing paired pre- and postassessment scores. "Overall" results contain all trainees. "Nonhealth" results include only the subset of trainees who indicate a "nonpublic health-related" pre-COVID-19 job classification, whereas "health" results include the subset of trainees who indicated having a pre-COVID-19 job as a disease investigator, in public health, or in an allied health field. Occasionally, trainees did not have career information available so are only included in the overall analysis; this applies to 14 trainees for the CT knowledge check, 15 for the CT self-assessment analysis, 102 for the CI knowledge check, and 95 for the CI self-assessment.

without requiring travel and to share lessons learned across local health jurisdictions. The success of the VTA should encourage the further development and use of online training programs, both during the COVID-19 pandemic and for future public health needs that arise. *AJPH*

ABOUT THE AUTHORS

Olivia Golston is with the Fielding School of Public Health, University of California, Los Angeles. Michael Prelep, Alina Dorian, and Shira Shafir are with the Department of Community Health Sciences, Fielding School of Public Health. Debbie Bain Brickley, Andrew Maher, Michael J. A. Reid, and Karen White are with the UCSF Pandemic Initiative for Equity and Action, University of California, San Francisco. Anne Cass is with the Center for Infectious Diseases, California Department of Public Health, San Diego. Lisa Chen is with the Curry International Tuberculosis Center,

University of California, San Francisco. Alice Gandelman is with the California Prevention Training Center, University of California, San Francisco. Chris Keh is with the Center for Infectious Diseases, California Department of Public Health, Richmond. Roger Myrick is with the Institute of Global Health Sciences, University of California, San Francisco. Rachel Willard-Grace is with the Center for Excellence in Primary Care, Department of Family and Community Medicine, University of California, San Francisco.

CORRESPONDENCE

Correspondence should be sent to Olivia Golston, Fielding School of Public Health, Department of Community Health Sciences, 650 Charles E. Young Dr South, 36-071 CHS, Box 951772, Los Angeles, CA 90095-1772 (e-mail: olivia.golston@ucla.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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O. Golston conducted the statistical analysis and drafted the article. O. Golston, M. Prelep, A. Cass, A. Dorian, R. Myrick, M. Reid, K. White, and S. Shafir conceptualized the analysis. M. Prelep, A. Cass, C. Keh, M. Reid, and S. Shafir reviewed and revised the article. D. B. Brickley, L. Chen, A. Gandelman, A. Maher, and R. Willard-Grace designed the Virtual Training Academy training program. All authors approved the final version of the article.

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HUMAN PARTICIPANT PROTECTION

The University of California, Los Angeles, Office of Human Research Protection Program determined that this project did not meet the definition of human participants research, so did not require institutional review board approval.

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Cannabis: Moving Forward, Protecting Health

Edited by: David H. Jernigan, PhD, Rebecca L. Ramirez MPH, Brian C. Castrucci, DrPH, Catherine D. Patterson, MPP, Grace Castillo, MPH

This new book addresses the ongoing debate on cannabis policy and provides guidance on how to regulate its sale and distribution. Instead of taking a stance for or against cannabis use, the book:

- suggests we employ strategies similar to those used in alcohol control to create a solid foundation of policy and best practices;
- focuses on how we can best regulate a complex substance.

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High-Risk Outreach for COVID-19 Mortality Reduction in an Indigenous Community

Myles J. Stone, MD, MPH, Ryan M. Close, MD, MPH, Christopher K. Jentoft, MD, Katherine Pocock, MHS, PA-C, Gwendena Lee-Gatewood, BS, Brooke I. Grow, Kristen H. Parker, PharmD, BCPS, April Twarkins, BSN, RN, J. T. Nashio, and James B. McAuley, MD, MPH

Indigenous populations have been disproportionately affected by COVID-19, particularly those in rural and remote locations. Their unique environments and risk factors demand an equally unique public health response. Our rural Native American community experienced one of the highest prevalence outbreaks in the world, and we developed an aggressive management strategy that appears to have had a considerable effect on mortality reduction. The results have implications far beyond pandemic response, and have reframed how our community addresses several complicated health challenges. (*Am J Public Health*. 2021; 111(11):1939–1941. <https://doi.org/10.2105/AJPH.2021.306472>)

More than 22% of our rural Native American community was diagnosed with COVID-19 between April 1, 2020 and February 15, 2021, an incidence nearly three times higher than that reported for the rest of the United States.¹ Medical care is provided by a single 22-bed hospital covering an area approximately the size of Delaware. The unique environment and scale of the outbreak necessitated an efficient and creative response to the pandemic.

INTERVENTION

We developed an early outreach and field medicine program to augment contact tracing and act as a bridge between community surveillance and hospital care.

PLACE AND TIME

Our program is based out of Whiteriver, Arizona, and serves approximately

18 000 people living on and around the White Mountain Apache Reservation. Outreach began in March 2020, and remains ongoing.

PERSON

Community members infected with or exposed to SARS-CoV-2 (severe acute respiratory syndrome coronavirus 2, the causative agent of COVID-19).

PURPOSE

The high transmissibility and evasive clinical presentation of SARS-CoV-2 often leads to patients being infected for days before they are traced and tested. Many people then enter medical care late in the disease process, when medical therapeutics are far less effective. Our goal was to quickly identify, monitor, and support those at highest risk of death.

IMPLEMENTATION

COVID-19 was first diagnosed in our community on April 1, 2020. Within weeks, our Indian Health Service facility was identifying more than 250 new cases per 100 000 persons daily. Quickly outpaced, we focused our efforts on finding the highest-risk contacts of each person and testing them immediately.

Our community's low phone ownership necessitated that most tracing activities be performed in the field. Testing was performed in contacts' homes at the time of tracing, which often meant that newly discovered cases were still in the asymptomatic period. As tracers returned to deliver test results or drop off isolation supplies, they noticed a concerning pattern among patients in the earliest stages of COVID-19. It was clear that the window of time between asymptomatic infection and critical illness was small for many, and particularly so for elders.

Short-interval follow-up was needed, but the usual methods of health care delivery would have overwhelmed local resources. Our facility was already operating near capacity, seeing about 1000 primary care encounters, 500 emergency department visits, and 50 admissions or transfers each week. We therefore partnered with tribal government to integrate public health outreach and direct medical care.

We built a High-Risk COVID Outreach Program to conduct repeated home visits for every person older than 60 years (and most older than 40 years) with known or suspected to have COVID-19. Hospital staff were paired with community health workers, and each team typically saw about 10 to 20 patients a day. Depending on the case volume, we would dispatch up to four teams each morning. The field physician would provide remote medical supervision for each team while also directly visiting patients with more complicated health needs. Visits for all patients continued until symptoms improved or the patient met criteria for hospital admission (most often ambulatory hypoxia).

Hospitalized patients would also receive a visit from the field team within 24 hours of discharge. This was critically important for patients who had been hospitalized at lower-elevation facilities. Blood oxygen levels often dropped considerably when patients returned home to our mountainous region. Patients could then be rehospitalized locally or set up with home oxygen and close follow-up depending on their risk of further decompensation. The latter option not only allowed patients to remain in a more comfortable environment, but also freed up scarce hospital beds for sicker patients.

EVALUATION

Integrating public health outreach and direct medical care appears to have had a considerable impact on our communitywide case fatality rate (CFR). It was 1.3% as of February 15, 2021, lower than Arizona's general population CFR of 1.9%, and just over a third of that for all other Native Americans statewide (3.2%).²

Our community had 3904 cases of COVID-19 by February 15, 2021. Of these, 1348 (34.5%) with a test date of January 15, 2021 or earlier were identified through contact tracing and followed in our program. The median age was 55 years (interquartile range = 41–64). This group is among those at highest risk of death from COVID-19,³ yet the CFR for all 1348 patients was 1.3%. This is lower than some reported general-population estimates.⁴

We wondered if we had simply identified more mild cases relative to other communities, but this does not seem to be the case. The COVID-19 hospitalization rate in our community was 24.7% ($n = 966$), markedly higher than that reported for the rest of the country (6.2%).¹ The CFR for patients in our program aged older than 60 years who were eventually hospitalized was 5.2% ($n = 10$). A major study reported a CFR of 25.7% in a general hospitalized population with similar age distribution.⁵ Within our own community, the CFR for hospitalized individuals aged older than 60 years who self-presented for medical care outside of the high-risk program was 14.3% ($n = 18$). The relative risk of death for those in the program was 0.36 (95% confidence interval = 0.17, 0.76; $P < .01$) despite having the same geographic distribution and

similar average age (70 vs 72 years for program participants and those self-presenting, respectively), average body mass index (31.1 vs 29.6 kg/m²), diabetes rate (58.1% vs 58.8%), and hypertension rate (70.1% vs 76.5%).

ADVERSE EFFECTS

No patients reported adverse effects. Staff members encountered impassable roads, wild animals, and the occasional patient who declined home visits (< 1%). By traveling during daylight hours in teams of at least two, the effects of these hazards were minimal.

SUSTAINABILITY

The partnership between our hospital, tribal leaders, and community members created a strong sense of unity and purpose that has strengthened other multidisciplinary partnerships addressing mental health and noncommunicable diseases unrelated to COVID-19. In fact, it led to the formation of a new Preventive Medicine Department that aims to apply the lessons learned from this program to other health conditions that have been challenging to address through usual care delivery models.

Staff members were initially recruited from departments that became less active during the pandemic. The full-time equivalents of one physician and one physician assistant were assigned to the program (of a total medical staff of about 30), along with one nurse and three medical assistants. Partial salary reimbursement came from billing for home visits and telehealth services, with additional investment from our hospital's operating budget to offset the difference. The benefits from

improved survival rates, reduced admissions, and enhanced community partnerships are difficult to quantify in purely financial terms, but we feel they are worth the modest cost of our program.

PUBLIC HEALTH SIGNIFICANCE

COVID-19 remains disproportionately hazardous to rural Indigenous communities. Our experience has shown that brief, frequent evaluations of people at risk for severe COVID-19 can extend the capacity of limited health care resources and have considerable impact on communitywide mortality. As a stand-alone intervention, this program had a better mortality risk reduction than commonly used medical treatments.⁵⁻⁷ But early outreach alone does not save lives, it simply allows therapeutics and health professionals to work more effectively when there is a scarcity of both.

The dichotomy between public health and medicine has always been artificial. Integrating the two saves lives, and our experience has been a striking reminder of this. As we rebuild our postpandemic health systems, the populations we serve will benefit from expanded coordination between public health agencies and traditional health care facilities. The shared goals have never been more apparent or more necessary. *AJPH*

ABOUT THE AUTHORS

Myles J. Stone and Kristen H. Parker are with the US Public Health Service Commissioned Corps, Rockville, MD. Ryan M. Close, Christopher K. Jentoft, Katherine Pocock, April Twarkins, and James B. McAuley are with the Whiteriver Indian Hospital, Whiteriver, AZ. Gwendena Lee-Gatewood and J. T. Nashio are with the White Mountain Apache Tribe, Whiteriver, AZ. Brooke I. Grow is with the Johns Hopkins Center for American Indian Health, Whiteriver, AZ.

CORRESPONDENCE

Correspondence should be sent to Myles Joseph Stone, 6300 N. Highway 89, Flagstaff, AZ 86004 (e-mail: myles.stone@ihs.gov). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

M.J. Stone was the director of the High-Risk Program and wrote the initial and revised manuscripts. R. M. Close oversaw the hospital-wide COVID-19 response, conceptualized the High-Risk Program, performed the statistical analysis, and contributed to subsequent iterations of the manuscript. C. K. Jentoft also served as a director of the High-Risk Program and, along with K. Pocock, performed chart reviews for the comparative analysis. G. Lee-Gatewood is the chairwoman of the White Mountain Apache Tribe and coordinated tribal support of this program and report. B. I. Grow performed the initial chart reviews for outcome analysis of program participants. K. H. Parker and A. Twarkins were field team leads and helped develop the described intervention approach. J. T. Nashio was the lead community health representative overseeing much of the field work by tribal staff. J. B. McAuley is the hospital clinical director and provided the staff and institutional support for the program. All authors contributed revision edits and approved the final draft.

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No authors have any potential or actual conflicts of interest to disclose.

HUMAN PARTICIPANT PROTECTION

This project was approved by the White Mountain Apache Tribal Council and Health Board and was

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The Economy as a Determinant of Pandemic Deaths

Kenneth Rochel de Camargo, MD, PhD

ABOUT THE AUTHOR

Kenneth Rochel de Camargo is with the Instituto de Medicina Social Hésio Cordeiro, Universidade do Estado do Rio de Janeiro and is an associate editor with AJPH.

See also Brenner, p. 1950, Pagán, p. 1947, and Nguyen et al., p. 1944.

The strong relationship between socioeconomic contexts and the health of individuals and populations has long been settled science and a core tenet of public health.¹ Measuring effects and examining possible causal pathways, however, is an ongoing challenge for researchers and policymakers alike.^{2,3}

Brenner's article (p. 1950) discusses, based on innovative methods, possible effects of the COVID-19–related economic downturn on economic and health indicators in the United States, focusing on excess deaths as a result of economic strife. Putting those effects into perspective, he also analyzes the health repercussions of the long recession that has afflicted the United States through much of the past decades, which also helps to demonstrate that some of the bad repercussions can happen in the short term.

The two accompanying editorials, by Pagán (p. 1947) and Nguyen et al. (p. 1944), both stress key issues raised by Brenner: (1) The pandemic had a clear negative effect on the economy that was responsible for an important amount of excess deaths. (2) The recession disproportionately affected communities of color and was independent of the implementation of social distancing measures. (3) Increases in the

health care workforce can attenuate some of the repercussions of the debilitated economy on the population's health.

The role of health care and prevention in softening the blow of economic hazards has been observed elsewhere; for instance, during one of the worst periods of the Brazilian economy throughout the 1980s, infant mortality, which has always been closely associated with general living conditions, kept steadily declining.⁴ This has been interpreted as a result of the expansion of the country's immunization program combined with specific measures addressing low-weight infants in the poorer regions. Community health workers tracked children's development to detect early signs of malnutrition, and oral rehydration therapy was implemented at the onset of cases of diarrhea.

The articles by Brenner, Pagán, and Nguyen et al. correctly point to the need to expand the public health workforce and provide wider access to health care as urgent measures to counter the broad negative repercussions of the pandemic, including its economic consequences. Nguyen et al. further emphasizes the need to incorporate research from other disciplinary fields, such as economics and labor

markets, to better understand causal relationships in the economic system as well as its intersections with people's health and possible pathways to alleviate the impact of economic downturns in socioeconomic inequalities and their repercussions on health.

There is, however, a limit to what public health interventions can achieve without addressing what many social epidemiologists have dubbed “the causes of causes,” that is, the overall structural social and economic issues that push entire populations into heightened vulnerability to poor health.

Economic programs that reduce wages and public spending to restore balanced state budgets⁵ have been shown to slow down economic recovery and further penalize the already fragile segments of the population.⁶ Discrimination based on race, gender, and sexual orientation further compounds the problem.

To achieve the long-sought aim of “health for all,” overarching political, economic, and social change is necessary to address the fundamental drivers of poor health and inequality.⁷ Without ensuring that everyone has adequate housing, good jobs, proper education, food security, and protection against all forms of discrimination—in summary, the bare minimum for a decent life—public health personnel will still be compensating for the damage and not building on fair foundations. **AJPH**

CORRESPONDENCE

Correspondence should be sent to Kenneth Camargo, Instituto de Medicina Social, Universidade do Estado do Rio de Janeiro, R. S. Fco. Xavier, 524 7o. andar Bloco D, Rio de Janeiro, RJ 25.559-900, Brazil (e-mail: kencamargo@gmail.com). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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COVID-19 Era Recessional Mortality Continues the Impact of Prepandemic Recessions on Mortality

Nga TQ. Nguyen, BPharm, PhD, Hoa Q. Nguyen, BPharm, PhD, and Ciaran O'Neill, PhD

ABOUT THE AUTHORS

Nga TQ. Nguyen is with the Department of Pharmaceutical Administration, University of Medicine and Pharmacy at Ho Chi Minh City, Vietnam. Hoa Q. Nguyen is with the Department of Clinical Pharmacy, University of Medicine and Pharmacy at Ho Chi Minh City, Vietnam and the School of Pharmacy, Queen's University, Belfast, UK. Ciaran O'Neill is with the Centre for Public Health, Queen's University, Belfast, UK.

 See also Brenner, p. 1950.

COVID-19 has had profound health, social, and economic effects globally. Its impact and our understanding of it will continue to evolve as issues—related, for example, to delayed diagnoses, social isolation, increased public sector borrowing, and health care worker burnout—emerge. Short-term effects, such as COVID-19–related deaths, are significant; in the United States these are estimated to have exceeded 630 000.¹ The longer-term effects will take time to evolve but are already sources of concern to decision makers and their consultants, including those in academia. In April 2020, the US unemployment rate reached 14.8%—the highest rate observed since data collection began in 1948.² In this issue of *AJPH*, Brenner (p. 1950) examines how this and other economic indicators of the effects of COVID-19 on the economy affect health. Brenner examines the relationship of unemployment and bankruptcy rates with excess deaths during the COVID-19 recession, contextualizing these with reference to

the Great Recession of 2000 to 2018 in the United States.

CONTINUING RECESSIONAL EFFECTS

Unemployment rates and gross domestic product (GDP) per capita have been long considered two major indicators of recession. Studies have shown rising unemployment and falling GDP per capita to be associated with higher mortality.^{3,4} Brenner used separate pooled cross-sectional time series analysis of state-aggregated mortality, per capita GDP, unemployment, and bankruptcy data for the Great Recession period (2000–2018) and the COVID-19 recession period (February–November 2020). The author found a strong positive relationship of unemployment, bankruptcy, and GDP per capita with excess deaths, controlling for COVID-19 cases, age, and race/ethnicity. Unemployment was also strongly correlated with major cause-specific mortality from 2000 to 2018. With these findings, it seems that the

recessional effect persists over the studied periods; that is, the COVID-19 era recessional mortality continues the consequences of prepandemic recessions (2000–2018) on multiple causes of mortality.

Brenner also examined the relationship between the recession and mental health disturbances. At the beginning of the COVID-19 pandemic, there were calls to focus on mental health services, as high rates of unemployment were predicted, potentially leading to increased suicide rates.⁵ Nevertheless, a recently published article has reported unchanged or decreased numbers of suicides across developed countries in the early months of the pandemic.⁶ These interesting findings could be because of early warnings of mental health issues during the pandemic, timely health care programs focusing on potentially susceptible people with both physical and mental illnesses, and fiscal policies enacted to support national economies.⁶ Equally, it could be that a rise in suicide is yet to emerge, mental health operating as a mediator.⁷ Indeed the relationship with mental health is complex. In a recent meta-analysis, Fond et al. reported a significant association between mental health disorders (e.g., major depression, anxiety, schizophrenia) and COVID-19–related mortality.⁷ Patients with mental health disorders may be prone to engage in behaviors (e.g., smoking, substance abuse) that made them more susceptible to COVID-19 compared with otherwise healthy people.

THE LONG-STANDING DISPARITY ISSUES

Communities left vulnerable by socioeconomic status, including Black and Hispanic communities as noted by

Brenner, could suffer disproportionately from both the Great Recession and the COVID-19 pandemic. This may be exacerbated in the case of COVID-19 because of these communities' greater representation as employees in personal services and hard-hit industries.⁸ A combination of psychological distress, inadequate access to health care insurance (arising from loss of employment), and social support during social distancing as well as insecurities in housing could place these communities in a relatively disadvantaged position for obtaining a job following the pandemic recession.

Systemic and mutually reinforcing sources of disadvantage were accentuated during both examined recessions and contributed to a widening gap between vulnerable communities and those who are better off in terms of their mortality experience. Brenner's study provides further evidence of this in the pandemic's impact on these communities. It also, though, highlights the positive contribution of health care in attenuating the mortality impact of both COVID-19 and broader economic shocks, offering hope and a warning.

WHERE DO WE GO FROM HERE?

As we noted, the effects of COVID-19 on health, health care, the economy, and society will take time to evolve. The effects may well be greater than those Brenner estimated, depending on how measurements are made, for example, whether we examine potential years of life lost rather than mortality and whether we include stillbirths in our analyses. Moreover, overall mortality picks up violence-related deaths, which may be an outcome of COVID-19-related recession and act as a mediator

in the relationship between unemployment and suicide. These are an important future area of research.

The moderating effect of the size of the health care workforce observed in Brenner's study and previous literature needs to be better understood as potentially related to countercyclical spending and the service actually delivered. This is in line with the perception that the health care system in the United States is "recession proof" based on economic evidence during the Great Recession.⁹ The fundamental principle for the inelasticity of health care demand in part is related to comprehensive insurance coverage with affordable deductible health plans.¹⁰ However, COVID-19 changed the situation and may lead to a shift from more generous private to public insurance as well as increased cost-averse behaviors by the privately insured.¹⁰ The COVID-19 crisis also led to a peak of 10.3% unemployment among those previously employed in the health care sector in April 2020.¹¹ Additionally, the increased stress level accentuated by fear of virus transmission and overwhelming workload because of COVID-19 also led to an estimated 49% of US health care workers experiencing burnout.¹² Therefore, timely investment to enlarge the health care workforce could be a priority to offset the adverse effects of COVID-19 on the health care system, which consequently contributes to save lives.

It is estimated that 10% unemployment is associated with approximately 48 149 excess deaths (35 700 excess deaths while being jointly analyzed with bankruptcies). Equipped with evidence such as that presented by Brenner, decision makers can plan to mitigate the effects. Importantly, Brenner highlights the importance of economists,

labor market specialists, and public health researchers working more closely to better understand causal relationships, the impact on inequalities, and how these inequalities might be mitigated. As Brenner says, these are potentially avoidable deaths requiring urgent action that prioritizes populations of low socioeconomic status and communities of color. **AJPH**

CORRESPONDENCE

Correspondence should be sent to Nga TQ. Nguyen, University of Medicine and Pharmacy at Ho Chi Minh City, 41-43 Dinh Tien Hoang, Ben Nghe Ward, District 1, Ho Chi Minh City Vietnam 71006 (e-mail: nguyenthiquynhnga@ump.edu.vn). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

N. TQ. Nguyen drafted the first version of the editorial. All authors revised, completed, and approved the editorial.

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Excess Deaths During the COVID-19 Economic Downturn

José A. Pagán, PhD

ABOUT THE AUTHOR

José A. Pagán is with the Department of Public Health Policy and Management, School of Global Public Health, New York University, New York, NY.

See also Brenner, p. 1950.

More than 700 000 people have died from COVID-19 in the United States over the last year and a half since the pandemic began.¹ Measures to stop the spread of COVID-19, such as stay-at-home orders and business closures, together with severe disruptions in overall economic activity brought about a short but deep economic recession. The Bureau of Economic Analysis of the US Department of Commerce estimates that real gross domestic product in the United States decreased at an annual rate of 19.2% from the fourth quarter of 2019 through the second quarter of 2020.²

In this issue of *AJPH*, Brenner (p. 1950) investigates excess deaths associated with unemployment and bankruptcies during the COVID-19 recession. The sharp economic downturn experienced by the United States last year is likely to lead to higher mortality over the next five to 10 years.³ The key question Brenner answers is whether higher unemployment and business closures due to the pandemic have a more immediate, same-year effect on excess deaths.

EXCESS DEATHS BEYOND COVID-19

Brenner uses data for the 50 US states from prepandemic years (2000–2018) and pandemic months (February 2020–November 2020), together with regression analysis, to study how the unemployment rate, GDP per capita with a 5-year lag, health care workforce size per 1000 total employment, and bankruptcies are related to excess deaths. His statistical approach controls for the number of COVID-19 cases and other key demographic and behavioral risk variables. The results of his study show that unemployment and bankruptcies during COVID-19 are significantly related to excess deaths. For example, 10% additional unemployment (i.e., unemployment rate as a proportion of the total labor force) is associated with 35 700 excess deaths, and a rate of 120 total bankruptcies per 100 000 people is associated with 144 483 excess deaths (Brenner, Table 2). Interestingly, his study finds that the size of the health care workforce also has a strong inverse relationship to excess deaths. This may have implications on the role of public health and health care delivery system

investments in mitigating the negative population health consequences of economic downturns.

There is no question that the COVID-19 pandemic has had a disproportionate impact on communities of color, particularly Blacks and Hispanics/Latinos. Survey data from April 2020 to March 2021 from the Understanding Coronavirus in America Tracking Study show that 89% of Latinos and 86% of Blacks experienced at least one serious economic, psychological, or health-related hardship compared with 76% of Whites.⁴ Brenner's study suggests that the impact of the pandemic will be even larger than expected for these two racial/ethnic groups given their high unemployment rates and their high presence in business sectors highly affected by the pandemic, such as the hospitality industry and retail and food services. In other words, not only are Blacks and Hispanics/Latinos at relatively high risk for COVID-19 because of where they live and work, but the population health and equity consequences of the pandemic are likely to be dramatic for them given the independent impact of unemployment and bankruptcies on excess deaths. What we are seeing now is just the tip of the iceberg in the sense that these economic effects take a few years to fully manifest themselves (i.e., they occur with a lag).

THE TIME TO ACT IS NOW

Brenner's study highlights the need to implement measures to address issues directly related to a specific event, such as a pandemic or a natural disaster. In addition, it points out that it is essential to immediately and swiftly confront

related challenges such as job loss and disruptions in economic activity in economic sectors and communities affected by a catastrophic event. Policy conversations often tend to focus on dealing with what may seem like the most critical, immediate needs (e.g., vaccine distribution in a pandemic, food distribution after a hurricane), but it is never too early to start the conversation and, more importantly, to act to meet other needs that become evident much later, when it could be too late (e.g., avoiding business closures or providing adequate unemployment assistance). The main consequence of not doing something in the face of a sharp economic downturn is a more pronounced impact on the most vulnerable. The experience from the COVID-19 recession and other economic downturns and natural disasters suggests that this will certainly be the case.

INVESTMENTS IN PUBLIC HEALTH INFRASTRUCTURE

An interesting finding of Brenner's study is that the size of the health care workforce seems to have partially suppressed the harmful effects of economic downturns during the 2000–2018 period. Although this finding may be difficult to interpret, it is consistent with the idea that a strong public health and health care infrastructure can help everyone to weather the negative effects of economic downturns on population health. The Affordable Care Act led to substantial investments in the health care infrastructure of the United States and expanded health insurance coverage.⁵ Perhaps these types of investments have had the unintended benefit of making it easier to address

COVID-19 challenges for public health and health care delivery systems. For example, the largest municipal health care system in the United States (New York City Health + Hospitals) reports that its ability to effectively serve COVID-19 patients, and all other patients, resulted from investments in an integrated electronic health system for its multiple safety net clinics and hospitals shortly before the pandemic began.⁶

THE FUTURE IS SOMEWHAT PREDICTABLE

Highly transmissible variants of the SARS-CoV-2 virus (the virus that causes COVID-19) have surfaced already and are likely to continue emerging without the adoption of strong mitigation measures.⁷ A more transmissible virus threatens to partially reverse the progress made to contain COVID-19 through vaccination and other preventive efforts. A virus resurgence and a subsequent economic slowdown will certainly have a disproportionate impact on the same populations that have experienced the worse effects of what feels like a never-ending pandemic.

Brenner's study provides important insights on what to expect in the immediate and near future in terms of mortality. The most recent life expectancy estimates from the National Vital Statistics System show that between 2019 and 2020, life expectancy at birth in the United States declined by 3.0 years for the Hispanic population, 2.9 years for the non-Hispanic Black population, and 1.2 years for the non-Hispanic White population.⁸ These sharp declines in life expectancy and the differences across racial/ethnic groups highlight the need to continue to delve deeper

into how recessions, pandemics, and natural disasters are related to population health and health disparities. *AJPH*

CORRESPONDENCE

Correspondence should be sent to José A. Pagán, Department of Public Health Policy and Management, School of Global Public Health, New York University, 708 Broadway, Room 719, New York, NY 10003 (e-mail: jose.pagan@nyu.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONFLICTS OF INTEREST

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Racism: Science & Tools for the Public Health Professional

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This important publication builds on the racial health equity work that public health advocates and others have been doing for decades. They have documented the existence of health inequities and have combatted health inequities stemming from racism. This book, which targets racism directly and includes the word squarely in its title, marks an important shift in the field's antiracism struggle for racial health equity. It is intended for use in a wide range of settings including health departments, schools, and in the private, public, and nonprofit sectors where public health professionals work.

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Unemployment, Bankruptcies, and Deaths From Multiple Causes in the COVID-19 Recession Compared With the 2000–2018 Great Recession Impact

M. Harvey Brenner, PhD

Objectives. To determine whether unemployment and bankruptcy rates are related to increased excess deaths during the COVID-19 recession and to examine whether the current recession-based mortality rate not only is dependent on COVID-19 but also continues the pattern of recessions, especially the Great Recession, in relation to chronic disease mortality rates and mental health disturbances (e.g., including suicide) from 2000 to 2018.

Methods. This study used pooled cross-sectional time series analysis to determine the impact of unemployment and bankruptcy rates on excess deaths from February to November 2020 for US states. The study used a second pooled cross-sectional time series analysis to determine whether the COVID-19-era recessional mortality continues the impact of prepandemic recessions (2000–2018) on multiple causes of mortality.

Results. Ten percent unemployment was associated with approximately 48 149 excess deaths, while, jointly with bankruptcies, their combined effect produced 35 700 and 144 483 excess deaths, for unemployment and bankruptcies, respectively. These health-damaging COVID-19-recession findings suggest a reiteration of the significantly increased major cause-specific mortality during 2000 to 2018, mitigated by the size of the health care workforce.

Conclusions. Minimization of deaths attributable to the COVID-19 recession requires ample funding for the unemployed and underemployed, especially Black and Hispanic communities, along with significant investments in the health workforce. (*Am J Public Health.* 2021;111(11):1950–1959. <https://doi.org/10.2105/AJPH.2021.306490>)

The United States continues to experience an unpredictable COVID-19 pandemic, during which deaths have been accelerating since November 2020, and the national toll has reached 4000 persons per day. Overall, deaths have exceeded 595 000 as of June 9, 2021. The Centers for Disease Control and Prevention has estimated a reduction, in the first half of 2020, in years of life expectancy—with

2.7 years lost by African Americans, 1.9 years lost by Hispanic populations, and 0.8 years lost by Whites.¹ The current COVID-19 pandemic is exacerbated by the appearance of newer variants originating in the United Kingdom, South Africa, and Brazil, which bring further uncertainty to the death rate. The loss of life expectancy raises the question whether—or to what degree—the estimated deaths are

partly the result of the abrupt national recession, which caused extraordinarily high unemployment rates and business closures in the onset of the COVID-19 pandemic. Is the accompanying recession a source of additionally increased excess deaths?² If the COVID-19-based recession has, in itself, produced higher-than-expected mortality, is this a unique feature of the COVID-19 pandemic—increased mortality would thus

be the result of a natural disaster—or is it a continuing effect of heightened unemployment and loss of income that, as evidenced during the Great Recession, may have increased national and state mortality during the first 2 decades of the 21st century?

The impact of COVID-19 on recession, especially unemployment, is clear from reports by economists.^{3,4} But does COVID-19–based unemployment by itself produce additional mortality beyond that initiated directly through the COVID-19 infection? Literature over 40 years, consistent with epidemiology at the individual level,^{5–9} shows, using a variety of methods, medium- to long-term (5- to 10-year) effects of recessions on elevated mortality (especially cardiovascular causes).^{10–15}

But are there short-term effects within the same year of increased unemployment and bankruptcies on excess deaths during the COVID-19 period of February to November 2020? And, if so, are these effects predictable resumptions of the impact of prepandemic recessions on mortality during the first 2 decades of the 21st century, which includes the Great Recession and its aftermath? The evidence so far shows conflicting findings.¹⁶ The present single-year analysis of the health damage of unemployment supports the relationship between recessional factors and mortality rates; it approximates the types of calculations typical of business cycle analysis (although unemployment is understood to be a “lagging” business cycle indicator).¹⁷

To answer these questions, I explored the potential impact of unemployment and bankruptcies on total excess deaths over the COVID-19 pandemic period of February to November 2020. I furthermore examined the relationship between employment status

and gross domestic product (GDP) declines during the prepandemic 2000–2018 period as potentially linked to mortality. Proximate causes of mortality include heart disease, cancer, stroke, diabetes, chronic lower respiratory disease, and suicide.

METHODS

In many observational studies, observations are available over a sequence of points in time (e.g., states and years as in our case). Examination of only 1 dimension (i.e., space or time) would limit us to perform classical cross-sectional or time-series regression analysis. Drawing on more advanced techniques^{18,19} (i.e., pooled cross-sectional time series analysis [PCSTS]), allows us to model simultaneously both space and time components as discussed by Reibling.²⁰

Pooled Cross-Sectional Time Series Analysis

The PCSTS method combines 2 approaches. First, the more familiar one is cross-sectional analysis, where, in this case, US states were the units of analysis. I examined multiple cross-sectional analyses corresponding to the 19-year period of 2000 to 2018 for which all of the data representing the individual variables were available with respect to all 50 US states. Second, the same is true for the PCSTS analysis of the monthly period of February to November 2020, which I analyzed by state. I based all variables used in these PCSTS analyses on aggregated data (i.e., population rates rather than individual-level data).²¹ In addition to the cross-sectionality of this procedure in both periods (2000–2018 and February–November 2020), the technique

simultaneously entails time-series analysis, involving variations over time in the individual predicted variables and the outcome variables—age-adjusted rates of overall mortality and those for major chronic diseases (apart from dementia) and suicide.²²

A more detailed discussion of the PCSTS approach is provided in Appendix A (available as a supplement to the online version of this article at <http://www.ajph.org>). References to the data sources for each of the variables are provided in Appendix B (available as a supplement to the online version of this article at <http://www.ajph.org>).

Major Variables in the Predictive Models

Three central variables, acting together, hypothetically promoted or harmed (in the case of the unemployment rate) overall population health during 2000 to 2018. These variables were the unemployment rate, GDP per capita with 5-year lag, and the size of the health care workforce. The lagged measure of GDP per capita was, hypothetically, the principal source of medium- and long-term national economic and population health gain.

Following GDP per capita with a 5-year lag, a key predictive variable was the very short-term impact of the unemployment rate, representing the immediacy of economic recession, during a single year (i.e., without lag). The third variable was the availability of health care represented by the size of the US health care workforce per 1000 of total employment. This variable is a potential central moderator of the effects of unemployment and GDP per capita with a 5-year lag. In this 3-variable overall hypothesis, all 3 variables act jointly to influence the overall

mortality rate. The study's argument required the modeling of all 3 variables simultaneously. Additional control variables (including severe poverty, tobacco, opioid and extensive alcohol consumption, and environmental pollution) are listed in detail in the section on control variables, including rationales, operationalization, and references for each control variable.

Components of the National Economy

Journalistically, the unemployment rate has been the principally used indicator measuring the presence of recessions. Despite alternative measures for which there are plausible arguments, this rate is nevertheless the metric that is most commonly used in policy discussions of the implications of recession, both short- and long-term.

While past researchers have typically either concentrated on the medium-term or lengthy effects of unemployment on damaged health, recent literature at the macroeconomic level has tended to focus on the near-simultaneous relation between unemployment and mortality. Such refocusing assumed that both unemployment and mortality could potentially behave like typical business cycle indicators, the effects of which are often seen within a single year. However, in the case of unemployment, it is understood by the National Bureau of Economic Research to be a lagging indicator, the maximum effect of which can remain high at least a year or 2 following the end of economic recessions.²³ Thus, to capture the health implications of unemployment within the same year that unemployment comes to its peak, it was necessary to treat the unemployment without any lag

(i.e., contemporaneously) from February to November 2020, thus capturing only very short-term effects. At the same time, I used estimates of GDP per capita with a 5-year lag to capture medium-term effects of recession and growth on mortality from 2000 to 2018.

Log Gross Domestic Product per Capita

The most comprehensive hypothesized variable benefiting population health is log GDP per capita with a 5-year lag for the period 2000 to 2018. This variable is as yet unavailable for February to November 2020. The GDP measure, with a 5-year lag, expresses the standard of living for the general society over the medium term within a full 5- to 6-year business cycle.²³ In the present study, I expressed GDP per capita in terms of a 5-year lag to take account of the cyclical implications of investments that require several years to result in health improvements (especially in health care technology and pharmaceuticals), in occupational–environmental health, and poverty minimization via social welfare expenditures affecting many categories of living conditions and elevated consumption of necessary goods and services (e.g., nutrition, clothing, shelter, electricity, transportation, rent).

Long-term GDP per capita has predominant importance for improved public health even though it includes elements of substantial economic inequality,²⁴ especially by US region, income group, educational status, and race/ethnicity. The economic inequality issue, in conjunction with long-term economic growth (for the general population and especially high-income groups) should not be underestimated.

In the United States, rural, Midwestern, and Rust-Belt areas have fallen behind in economic growth and life span. It has been argued that these aspects of lagging development have been at least partly responsible for the emergent trends of the opioid crisis among younger populations and long-term loss in life expectancy (i.e., “deaths of despair”).²⁵

Health care workforce (2000–2018). The size of the health care workforce over long periods in US history reflects a sustained upward trend. At the same time, the short-term and long-term effects of recessions have decreased the size of the health care workforce because of a loss of health insurance related to unemployment during recessions and a recession loss of income to the population more generally, thus inhibiting investment in an expanded workforce. Thus, I separately investigated the trend in the size of the health care system, highly contingent on scientific and technological advances and long-term health policy, as to its long-term implications for societal health.

Total bankruptcies (February–November 2020). In the February to November 2020 analysis, in addition to unemployment, I also used bankruptcies (private and commercial) as a COVID-19 recession predictor of excess deaths. In the 2020 monthly analyses of the impact of recession, it was important to take into account a somewhat broader set of measures—especially those that influence business as a whole. A prominent and traditional business cycle indicator is total bankruptcies,²⁶ which provides a more widespread sense of the extensiveness and depth of the COVID-19–initiated recession. The total bankruptcy rate

provides an indication of losses to small and large businesses.

Control variables. I introduced behavioral risk factors as variables into the predictive models to distinguish them from the more direct effects of income change or unemployment.

The behavioral risk factors from 2000 to 2018 were as follows:

1. Major depressive episodes: The intention was to discriminate between depressive episodes that were clearly linked to recession and those that were not necessarily associated with recession, but rather emanated from major life events²⁷ and daily hassles,²⁸ including those that occurred at the workplace, in family life, and especially resulting from the loss and grief related to damaged health and mortality. Nevertheless, depression is a significant risk factor for poor health, low life satisfaction, and early mortality.
2. Smoking: Smoking, a basic behavioral risk factor, was measured at a 3-year lag because the long-term trend in industrialized societies (especially the United States) has involved major declines in smoking prevalence, greatly curtailing mortality from cardiovascular illnesses, malignancies, and chronic obstructive pulmonary disease. The 3-year lag was hypothesized because there is evidence that several years are often required after smoking cessation for the former smokers' health to improve to a point that it returns to previous nonsmoker cardiopulmonary functioning.
3. Participation in physical activities: It is now widely accepted that participation in physical activities is a major source of health maintenance and improvement.
4. Alcohol-induced death rate: The literature is complicated regarding the health impact of the overall population consumption of alcohol. Especially for the cardiovascular illnesses, very high as well as virtually zero consumption are associated with elevated mortality, whereas moderate consumption appears to enhance longevity.²⁹ To specifically indicate chronically higher alcohol consumption, which elevates mortality for many chronic and mental health causes, I used age-adjusted alcohol-related deaths as a behavioral risk factor for mortality.
5. Tuberculosis: Tuberculosis incidence is especially high in low-income developing countries, but modest in industrialized countries. I used it as a predictive behavioral risk factor because of its intense association with chronic poverty, but not necessarily with trends in GDP or recession.

For references and further discussion of control variables, see Appendix C (available as a supplement to the online version of this article at <http://www.ajph.org>).

RESULTS

In the COVID-19 period of February to November 2020, unemployment was a significant predictor of excess deaths controlling for the number of COVID-19 cases, age, and Black and Hispanic racial/ethnic groups (Table 1). Using the same model, with identical controls, the combination of both recessional factors of unemployment and bankruptcies yielded an impact of further increased excess deaths (Table 2). In the analysis presented in Table 2, which combines

unemployment and bankruptcies, the numerical effect on excess deaths related to unemployment slightly decreased because of the relationship between unemployment and bankruptcies that typically would occur in a recession. In the model that presents 10% unemployment (as currently estimated by the Federal Reserve and Department of the Treasury³⁰) as the only recessional variable, an additional 48 149 deaths were estimated (Table 1). In the recessional model that included both unemployment and bankruptcies, the estimates of 10% additional unemployment led to 35 700 excess deaths, and a 120-unit increase per 100 000 in bankruptcies led to approximately 144 483 deaths.

Is this finding a unique, natural consequence of a recession related to a world pandemic or a partial reinstatement of the effects of recessions during the 21st century, specifically 2000 to 2018? Like total mortality, all major chronic disease causes of death showed significant beneficial effects of log GDP per capita with a 5-year lag and mortality increases related to unemployment without lag (Table 3).

The principal recessional variable, unemployment, showed evidence of elevated mortality during 2000 to 2018 (total mortality) and major chronic causes of death and suicide and continued to do so in the period February to November 2020 for total excess death. This indicates a continuous relationship of recession (especially indicated by unemployment rates) through the 2 research sample periods elevating mortality. The principal control variables showed strong positive relations to total mortality and mortality by major cause. However, the health care workforce with inverse relations to mortality featured as the strongest coefficient

TABLE 1— Cumulated Excess Deaths Rate per 100 000, All Causes, Pooled Cross-Sectional Time Series Analysis: 50 US States, February–November 2020

Variables	Cumulated Excess Deaths, All Causes ^a	Coefficient (P)
Cumulated incident cases rate/100 000		0.02 (<.001)
2000-unit increase	150 537	
4000-unit increase	301 074	
6000-unit increase	451 612	
8000-unit increase	602 149	
10 000-unit increase	752 686	
Unemployment rate as proportion (%) of total labor force aged ≥ 16 years, seasonally adjusted		1.47 (<.001)
5-unit increase	24 074	
10-unit increase	48 149	
15-unit increase	72 223	
20-unit increase	96 297	
25-unit increase	120 372	
Percentage of population aged ≥ 75 years, July 1, 2019, estimate		7.62 (.018)
1-unit increase	24 997	
2-unit increase	49 994	
3-unit increase	74 991	
4-unit increase	99 988	
5-unit increase	124 985	
Percentage of Black or African not Hispanic or Latino, July 1, 2019, estimate		1.03 (.001)
5-unit increase	16 927	
10-unit increase	33 855	
20-unit increase	67 709	
30-unit increase	101 564	
40-unit increase	135 418	
Percentage of White Hispanic or Latino, July 1, 2019, estimate		0.28 (.35)
5-unit increase	4517	
10-unit increase	9034	
20-unit increase	18 067	
30-unit increase	27 101	
40-unit increase	36 135	
Constant		−59.78 (.014)

Note. Data show expected additional excess deaths, all causes, when the independent variable increases by the specified units. The estimated coefficients from the excess deaths model were used to calculate the associated change of the cumulated excess deaths rate, all causes. Based on the total US population from the year 2019, the change of the cumulated excess deaths rate was converted to expected additional excess deaths numbers (e.g., a change of the unemployment rate by absolute 10% resulted in 48 149 additional excess deaths). The potential changes of the independent variables were derived by inspection of the descriptive statistics for each variable. Data and their sources of control variables are listed in Appendix B (available as a supplement to the online version of this article at <http://www.ajph.org>). For all key variables, a summary of descriptive statistics is provided in Appendix E (available as a supplement to the online version of this article at <http://www.ajph.org>).

^aEstimates for excess deaths were based on a 2019 US population of 328 239 523. The number of observations for the analysis was $n = 500$; $R^2 = .65$; Wald $\chi^2(5)$ (Prob > χ^2) = 1434.06 ($P < .001$).

among predictive variables except GDP per capita with a 5-year lag (Table 3).

The impact of recessions during 2000 to 2018 on suicide was clear, while the

predictive model and, therefore, the findings, were somewhat different from those of total mortality and that for major chronic diseases. For suicide,

the unemployment rate proved to be a robust and significant related predictive factor. GDP per capita, the major indicator of national economic change,

TABLE 2— Cumulated Excess Deaths Rate per 100 000, All Causes, Pooled Cross-Sectional Time Series Analysis: 50 US States, February–November 2020

Variables	Cumulated Excess Deaths, All Causes ^a	Coefficient (P)
Cumulated incident cases rate/100 000		0.01 (< .001)
2000-unit increase	98 141	
4000-unit increase	196 282	
6000-unit increase	294 423	
8000-unit increase	392 564	
10 000-unit increase	490 705	
Unemployment rate as proportion (%) of total labor force aged ≥ 16 years, seasonally adjusted		1.09 (< .001)
5-unit increase	17 850	
10-unit increase	35 700	
15-unit increase	53 549	
20-unit increase	71 399	
25-unit increase	89 249	
Cumulated total bankruptcies rate/100 000		0.37 (< .001)
60-unit increase	72 242	
120-unit increase	144 483	
180-unit increase	216 725	
240-unit increase	288 967	
300-unit increase	361 208	
Percentage of population aged ≥ 75 years, July 1, 2019, estimate		7.99 (.015)
1-unit increase	26 243	
2-unit increase	52 485	
3-unit increase	78 728	
4-unit increase	104 971	
5-unit increase	131 213	
Percentage of Black or African not Hispanic or Latino, July 1, 2019, estimate		0.48 (.14)
5-unit increase	7 911	
10-unit increase	15 822	
20-unit increase	31 643	
30-unit increase	47 465	
40-unit increase	63 286	
Percentage of White Hispanic or Latino, July 1, 2019, estimate		0.42 (.16)
5-unit increase	6 846	
10-unit increase	13 692	
20-unit increase	27 385	
30-unit increase	41 077	
40-unit increase	54 770	
Constant		−73.19 (.003)

Note. Data show expected additional excess deaths, all causes, when the independent variable increases by the specified units. The estimated coefficients from the excess deaths model were used to calculate the associated change of the cumulated excess deaths rate, all causes. Based on the total US population from the year 2019, the change of the cumulated excess death rate was converted to expected additional excess deaths numbers (e.g., a change of the unemployment rate by absolute 10% results in 35 700 additional excess deaths). Analogously, when the cumulated bankruptcies rate per 100 000 changes by 120 units, then this will result in 144 483 additional excess deaths. The potential changes of the independent variables were derived by inspection of the descriptive statistics for each variable. Data and their sources of control variables are listed in Appendix B (available as a supplement to the online version of this article at <http://www.ajph.org>). For all key variables, a summary of descriptive statistics is provided in Appendix E (available as a supplement to the online version of this article at <http://www.ajph.org>).

^aEstimates for excess deaths were based on a 2019 US population of 328 239 523. The number of observations for the analysis was $n = 500$; $R^2 = .62$; Wald $\chi^2(5)$ (Prob > χ^2) = 2053.66 ($P < .001$).

TABLE 3— Six Causes of Death, Rates per 100 000, Pooled Cross-Sectional Time-Series Models: 50 US States, 2000–2018

Variables	Total (ICD-10: A00–Y89), Coefficient (P)	Cancer (ICD-10: C00–D48), Coefficient (P)	Diabetes (ICD-10: E10–E14), Coefficient (P)	Heart Diseases (ICD-10: I05–I52), Coefficient (P)	Stroke (ICD-10: I60–I69), Coefficient (P)	CLRD (ICD-10: J40–J47), Coefficient (P)
Log normal of 5-y lag of real GDP in thousands of chained 2012 \$ per capita	−173.66 (< .001)	−43.02 (< .001)	−10.65 (< .001)	−116.02 (< .001)	−38.19 (< .001)	−5.65 (< .001)
Unemployment rate as proportion (%) of total labor force aged ≥ 16 y	1.72 (< .001)	1.70 (< .001)	0.14 (< .006)	1.06 (< .001)	0.18 (.052)	0.26 (< .001)
Employment in health care support per 1000 of total employment	−35.69 (< .001)	−8.07 (< .001)	−2.00 (< .001)	−20.26 (< .001)	−5.24 (< .001)	−0.77 (.12)
Tuberculosis incidence rate per 100 000 of total population	12.69 (< .001)	3.21 (< .001)	0.26 (.005)	8.62 (< .001)	2.06 (< .001)	−0.71 (< .001)
Major depressive episode in the past year (%) in total population aged ≥ 18 y	7.25 (< .001)	1.11 (< .001)	0.62 (< .001)	4.91 (< .001)	1.31 (< .001)	−0.40 (.013)
5-y lag of daily smoker prevalence in total population aged ≥ 18 y	5.85 (< .001)	2.50 (< .001)	0.15 (.002)	2.90 (< .001)	0.77 (< .001)	0.54 (< .001)
Age-adjusted alcohol-induced death rate per 100 000 of total population	1.03 (.01)	−0.97 (< .001)	0.19 (< .001)	−0.35 (.12)	−0.39 (< .001)	0.11 (.08)
Prevalence of not participated in any physical activities in population aged ≥ 18 y	2.46 (< .001)	0.09 (.29)	−0.01 (.78)	0.94 (< .001)	0.07 (.27)	0.4 (< .001)
Constant	1313.05 (< .001)	300.19 (< .001)	60.11 (< .001)	554.51 (< .001)	177.65 (< .001)	54.29 (< .001)
R ² overall	0.86	0.89	0.63	0.85	0.78	0.75
Wald χ^2 (23) (Prob > χ^2)	5504.14 (< .001)	6102.29 (< .001)	676.04 (< .001)	6488.50 (< .001)	4133.83 (< .001)	452.07 (< .001)

Note. CLRD = chronic lower respiratory diseases; GDP = gross domestic product; ICD-10 = *International Classification of Diseases, 10th Revision* (Geneva, Switzerland: World Health Organization; 1992). The number of observations was 950. All models have been adjusted with 15 regional dummies for New England, Mideast, Great Lakes, Southwest, Rocky Mountains, Far West, Alaska, Arizona, California, Florida, Hawaii, Louisiana, Nevada, Oklahoma, and Wyoming. Data and their sources of control variables are listed in Appendix B (available as a supplement to the online version of this article at <http://www.ajph.org>). For a detailed discussion of each of the control variables, see Appendix C (available as a supplement to the online version of this article at <http://www.ajph.org>). For all key variables, a summary of descriptive statistics is provided in Appendix E (available as a supplement to the online version of this article at <http://www.ajph.org>).

was not predictive for suicide, while, when instead median household income for GDP with 5-year lag (which de-emphasizes income inequality) was included, the model showed the hypothesized relationship. However, the control variables, expressing stress relationships and coping mechanisms, were positively related to and constituted significant risk factors in the occurrence of suicide (Table 4).

DISCUSSION

In the COVID-19 recession, both unemployment and bankruptcies exerted a substantial damaging impact on excess deaths. COVID-19 incidence, age, and race/ethnicity were controlled. This means that the ethnic/racial factors, which have been widely reported as being especially important,³¹ were

adjusted for in these 2 models covering the COVID-19 recession era. In these models, the potential importance of race/ethnicity in the COVID-19 recession was further highlighted by the fact that unemployment rates in 2020 were considerably higher for Hispanic and Black populations than for the White population³² and, for the same period, bankruptcies were especially high for

TABLE 4— Suicide Death Rate per 100 000, Pooled Cross-Sectional Time Series Model: 50 US States, 2000–2018

Variables	Suicide (ICD-10: U03, X60–X84, Y87.0), Coefficient (P)
Log normal of 5-y lag of median household income in chained 2018 \$	–1.28 (.043)
Unemployment rate as % of total labor force aged ≥ 16 y	0.09 (<.001)
Illicit drug use in the past month (%) in population aged ≥ 26 y	0.14 (<.001)
Gallons of alcohol per capita aged 14 y and older	1.87 (<.001)
Age-adjusted alcohol-induced death rate per 100 000 of total population	0.38 (<.001)
Age-adjusted nontransport accidents (W00–X59, Y86) death rate per 100 000 of total population	0.08 (<.001)
Nonmedical use of pain relievers in past year (%) in population aged ≥ 26 y	0.11 (.1)
Major depressive episode in the past year (%) in population aged ≥ 18 y	0.16 (.007)
Constant	15.37 (.024)

Note. ICD-10 = *International Classification of Diseases, 10th Revision* (Geneva, Switzerland: World Health Organization; 1992). The model has been adjusted with 6 regional dummies for New England, Mideast, Great Lakes, Rocky Mountains, Alaska, and California. The number of observations was 950; overall $R^2 = .79$; Wald χ^2 (23) (Prob > χ^2) = 1936.32 ($P < .001$). Data and their sources of control variables are listed in Appendix B (available as a supplement to the online version of this article at <http://www.ajph.org>). For a detailed discussion of each of the control variables, see Appendix C (available as a supplement to the online version of this article at <http://www.ajph.org>). For all key variables, a summary of descriptive statistics is provided in Appendix E (available as a supplement to the online version of this article at <http://www.ajph.org>).

the African American population, especially Black women.³³

However, the Federal Reserve and Department of the Treasury announced on February 22, 2021, that the official unemployment rate of 3.6% should really be taken as close to 10.0% because of many unemployed persons completely leaving the labor force.³⁰ If that is correct, then our estimates of the impact of unemployment could be underestimated.

Comorbidities of COVID-19 Mortality

In addition, the recessional effects of COVID-19 could also extend to the chronic disease comorbidities of COVID-19 mortality, as was the case for elevated chronic disease mortality in the 2000–2018 period. Main contributing factors in the COVID-19 pandemic could include the influence of psychological stress, loss of access to health care because of pressure on the health care system, and loss of

financial resources. In such cases, because the Black population has been more susceptible to many chronic diseases, their relatively low economic and occupational status may well put them at greater risk for recessional losses. The important implication is that far more extensive work is required to understand how key elements of economic recession intersect with race/ethnicity to produce much higher-than-expected mortality rates.

Support for the long-term relationship between unemployment and increased mortality has been found at the national level^{9,12–15} and extensively tested for at the individual level in epidemiological studies over 40 years.^{5–8,10,11} A principal contribution of this article is the demonstration that short-term, intense, and abrupt increases in unemployment have led to elevated excess deaths during the COVID-19 recession as well as expanded total and chronic disease mortality during the 2000–2018 period. Despite the fact that the

unemployment-to-mortality relationship has been established for medium- and long-term relationships,^{5,6,14} it is only recently that economists have begun to study this relationship at the macro level within a single year (i.e., the same year) in the attempt to use temporally coincident (i.e., contemporaneous) indicators in the technical business cycle terminology.

At first sight, it would seem that examining only the very-short-term mortality impact of unemployment would be counterintuitive with respect to the epidemiological tradition of research, which stipulates that the development of chronic diseases occurs over much of the life span. Previous research using this short-term approach has produced contradictory findings.¹⁶ Nevertheless, in keeping with the style of research that is specific to short-term business cycle analysis, I analyzed the measures of the effects of recession, unemployment over 2020 and 2000 to 2018, and bankruptcy rates during 2020 (over a single year) in

this article. At the same time, longer-term analysis by economists have emphasized the effects, after a first year of peak unemployment, of hysteresis, or scarring.³⁴ Hysteresis refers to the multiple effects lagging high recessionary unemployment during which employment and income losses persist.

Nevertheless, the mortality-inducing effects of long-term and severe economic loss are supported by the most prominent and ubiquitous findings in epidemiology, in industrialized countries and globally. This finding has become known as the “social gradient” or “health gradient.”³⁵ It specifies that the higher the level of socioeconomic status of a person or population, the lower will be the mortality rate. The direct inference is that elevation of socioeconomic status decreases population and individual mortality, while declines in socioeconomic status produce increases in mortality. This study is apparently one of the first that deals with the importance of GDP per capita, on both an annual and long-term basis, in the reduction of mortality for major causes during 2000 to 2018. In this article, the value added is that statistical tests were used to examine this hypothesis on a US state basis. The analysis over 2000 to 2018 provides the extensive statistical degrees of freedom necessary to examine the impact of the size of the health care workforce on multiple causes of mortality, holding constant overall GDP per capita. The resulting significant impact of the health care workforce is one of the major sources of mortality reduction during 2000 to 2018. It is clear that investment in the health workforce, including public health, has been a major factor leading to improvement in the public’s health. Indeed, it is possible that, absent the mortality-reducing

effects of the health care workforce during 2000 to 2018, the impacts of recession on increased mortality may have been greatly enlarged.

Public Health Implications

The COVID-19–induced unemployment and bankruptcy rates are robustly related to increased excess mortality from February to November 2020. Economists and labor market specialists should therefore join public health researchers to formulate policies that reduce mortality. Rapid policy intervention is especially required for populations of low socioeconomic status and communities of color who have suffered inordinately from COVID-19 in terms of morbidity and mortality. The vulnerability of these groups is attributable not only to occupations that more frequently involve interpersonal contact but also to low socioeconomic position, with considerably higher unemployment, income loss, and bankruptcy occurring to these populations in recession. We need to reconsider the epidemiology–public health disciplines involving natural disasters—epidemics, heatwaves, floods, hurricanes, etc. The typical approach in public health disaster relief has been to concentrate on mitigating the immediate disaster. But the corollary economic and social implications of natural disasters could have medium- and long-term implications for substantially expanded illness and mortality if these corollary effects are not attended to.

Although the COVID-19 recession is unique in being caused by a natural disaster, it follows a pattern of recessions during the 21st century, including the Great Recession, of increasing the total mortality rate and mortality specifically attributable to major chronic

diseases, such as heart disease, cancer, stroke, diabetes, and chronic lower respiratory disease, as well as mental health consequences, such as suicide.³⁶ It is clearly important for Congress to provide sufficient funds to aid in vaccine distribution and COVID-19 treatment of the full population. At the same time, the minimization of deaths attributable to the COVID-19 recession requires ample funding for the unemployed and underemployed, and to individuals and businesses who have experienced and are experiencing major economic losses. Also, it is especially important for funding to mitigate the maximum effect of recession to communities of color, rural populations, and those “left behind” in previous governmental support efforts during the Great Recession era. Given the vulnerabilities of communities of color to the pandemic and its economic consequences, it would be important to investigate more precisely how different racial/ethnic groups’ health vulnerabilities interact with economic losses imposed by recessions of the 21st century. Furthermore, the pre-pandemic 2000–2018 analysis makes clear that the absolute size of the health workforce is a highly significant factor in mitigating the results of economic recession, low socioeconomic status, and poverty. Investments into enlarging the size of the health workforce, and more generally into public health, is an important priority for reducing overall health inequalities in American society, in both the short and long term. **AJPH**

ABOUT THE AUTHOR

M. Harvey Brenner is with the Department of Health Policy and Management, Bloomberg School of Public Health, Johns Hopkins University, Baltimore, MD. He is also with the Department of Health Behavior and Health Systems, School of Public Health, University of North Texas Health Science Center, Fort Worth.

CORRESPONDENCE

Correspondence should be sent to M. Harvey Brenner, Johns Hopkins University Bloomberg School of Public Health, Department of Health Policy and Management, 624 N Broadway, Room 490, Baltimore, MD 21205 (e-mail: mbrenne2@jhu.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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The Social Survey, the Metropolitan Life Insurance Company, and the Beginnings of the US Public Health Service's Sickness Surveys

Mónica García, MD, PhD

The earliest sickness survey of the US Public Health Service, which started in 1915, was the Service's first socioeconomic study of an industrial community. It was also the first to define illness as a person's inability to work. The survey incorporated the Metropolitan Life Insurance Company's definition of illness, which, instead of sickness rates, focused on duration of illness as a proxy of time lost from work. This kind of survey took place in the broader context of the reform movements of the Progressive Era and the social surveys conducted in the United States, which led to the creation of the Federal Commission on Industrial Relations, where the Service's sickness survey originated. The Service's focus on the socioeconomic classification of families and definition of illness as the inability to work enabled it to show the strong link between poverty and illness among industrial workers. The leader of the survey, Edgar Sydenstricker, and the Metropolitan Life Insurance Company came up with new ways to measure the health of the population, which also influenced the Service's studies of the effects of the Great Depression on public health and the National Health Survey of 1935–1936. (*Am J Public Health*. 2021; 111(11):1960–1968. <https://doi.org/10.2105/AJPH.2021.306454>)

This is a study of the pioneering socioeconomic sickness survey in the United States, undertaken by the US Public Health Service (hereafter referred to as "Service") in the 1910s. The Service was created in 1912 to assist the public health work of the states. Between 1912 and World War II, the Service carried out socioeconomic investigations, which had an enormous influence on the scientific and public understanding of inequalities in health.¹ These investigations also provided the basis for innovative epidemiological designs.² The Service's early socioeconomic studies were also an adaptation of the social surveys that expanded after Charles Booth's work on poverty, *Life and Labor of the People in London*

(1899 and 1903), and Benjamin Seebohm Rowntree's study of poverty in York, England, *A Study of Town Life* (1901).³ The social survey in England studied the urban working classes to formulate programs to ameliorate the harsh conditions of their lives.⁴ The term refers to a number of investigations that, contrary to previous ones, entailed the first-hand collection of data about individuals, families, and households on a local rather than a national level and developed in tandem with the evolution of public policy and social reform.⁵ Among the most notable were those undertaken by the settlement house movement in England and the United States (i.e., the *Hull House Maps and Papers* about Chicago, Illinois,

1890), the work of W. E. B. Du Bois (1889), and the surveys sponsored by the Russell Sage Foundation (the *Pittsburgh Survey*, 1907–1908).⁶ However, historians have largely ignored the social surveys devoted to public health. The bibliography of social surveys by Allan Eaton and Shelby M. Harrison of the Russell Sage Foundation (1930) shows that public health was indeed at the core of the social survey.⁷ By focusing on the work of the first economist-statistician of the Service, Edgar Sydenstricker,⁸ I will show how the Service's investigations were a continuance of the Victorian and Progressive tradition.

One aim of the social surveys in the United States was to be objective and scientific, which, in the period in

question, was synonymous with quantitative methods. Indeed, the Progressive reformers increasingly relied on statistics, economics, and the emerging social sciences: the Service's early surveys adopted the quantitative analyses typical of the social survey.⁹ Of course, neither quantification nor the effort to measure the relation between social conditions and health was a novelty at that time. In the 19th century, public health advanced in close connection with the development of statistics and the rationalization of government administration. Starting in 1830, a French doctor, Louis René Villermé, pioneered quantitative investigations that established links between poverty and the poor health of the working classes in Paris, France,¹⁰ while in England, the General Register Office, created in 1837, developed an institutional structure for the study of the relationship between mortality, causes of death, and occupation that became the basis of the British public health movement.¹¹ In the United States, during roughly the same period, the decentralization of its government and the lack of a comprehensive compilation of vital statistics meant that there were no systematic studies of mortality patterns in the whole population before the 1930s. However, as Krieger and Fee have pointed out, investigations of the links between socioeconomic conditions and health by federal agencies and nongovernmental bodies were crucial in revealing the socioeconomic factors responsible for inequalities in health in the first half of the 20th century.¹²

One novel feature of the Service's surveys, led by Sydenstricker, which quantified the relation between socioeconomic conditions and health, was the employment of the methods pioneered by the Metropolitan Life

Insurance Company. As some historians have noted, the insurance companies were responsible for some of the most thorough mortality studies in that period.¹³ In Britain, the statistician William Farr of the General Register Office used the life tables technique of actuaries to calculate life expectancy, age-specific mortality rates, and standardized mortality rates to determine which districts of the country were healthy or not and improve the conditions of public health, which became the basis of the 20th century British classification of health conditions in terms of occupation.¹⁴ To throw further light on the links between US life insurance companies and public health surveys, this article shows how the Service's investigations implemented the measurements the Metropolitan Life Insurance Company used to investigate the health of its policyholders. The company's approach not only influenced the Service's earliest survey but also continued to be used up to the 1930s.

SOCIAL SURVEY AND THE SERVICE

Soon after its creation, the Service came up with a research agenda that would eventually make poverty the main focus of its analyses of the effect of socioeconomic differences on health.¹⁵ Sydenstricker was a graduate of Washington and Lee University (1902) and Fellow of Political Economy at the University of Chicago (1907–1908).¹⁶ His previous work in labor economics for the Immigration Commission (1907–1909)¹⁷ but, most importantly, for the Commission on Industrial Relations (1912–1915), connects Sydenstricker's work in the Service with the social survey movement initiated in Britain.

The term social survey was derived from Booth's work. Alarmed by the situation of the poor, certain sectors of the middle and upper classes drafted policies for social reform based on a systematic accumulation of data.¹⁸ Quantification was a significant aspect of the approach of Booth and his followers. Rowntree's work on poverty in York (1901), for example, sought "the true measure of the poverty in the city... how much of it was due to insufficiency of income and how much to improvidence."¹⁹ Booth specifically set out to show "the numerical relation which poverty, misery, and depravity bear to regular earnings and comparative comfort" and measure the number of people above and below the poverty line.²⁰ Instead of relying on official statistics, Booth innovatively compiled first-hand data about the household circumstances of poor families gathered by school board visitors. His findings not only had a direct impact on public policies and led to movements for social reform in England, but they also inspired pioneering social researchers in the United States, including Jane Addams and W. E. B. Du Bois.²¹

The social survey flourished in the United States in the early decades of the 20th century, thanks to the work of the philanthropist Margaret Olivia Sage and the magazine *Charities and the Commons* (renamed *Survey* in 1909), a leading national outlet for the reform movement.²² The magazine's editors, Edward T. Devine and Paul U. Kellogg, with financial support from the Russell Sage Foundation, undertook the Pittsburgh Survey in 1907–1908, which studied how industrial capitalism (the steel industry) shaped urban development.²³ After the bombing of the *Los Angeles Times* building in 1910 and the subsequent conviction and

imprisonment of those held responsible for it— John J. McNamara (the secretary of the International Association of Bridge and Structural Iron Workers) and his brother, James McNamara— these editors called on President Taft to create a commission for the investigation of social unrest. The bombing was the culmination of two decades of conflict between capital and labor caused by demands for better working conditions and the right to form trade unions. Along with leaders of the social survey movement, like Addams, the editors argued that the growing power of capitalism in American life was creating unrest in the working class, and because this problem was beyond the capacity of the country's legal system, a thorough investigation of trade unions and the cost of strikes would be needed to deal with it.²⁴

Thus, pressured by the Progressive reformers and leaders of the US social survey movement, the US Congress created the Commission on Industrial Relations in 1912 to ascertain the causes of industrial strife and ensure decent working and living conditions.²⁵ William Jett Lauck,²⁶ who had studied industrial communities for the Immigration Commission of 1907–1909,²⁷ was appointed director of research. Sydenstricker assisted the sanitary adviser of the Commission on Industrial Relations, the Surgeon General of the Service, Benjamin S. Warren, in the investigation of health insurance for wage earners and the living conditions of industrial workers and communities.²⁸

A staff of more than 50 persons was assembled for the Commission, most of them young economists.²⁹ Chiefly concerned with the strikes and violent protests that took place between 1910 and 1915, they interviewed more than 700 persons all over the country,

from businessmen to miners, farmers, garment workers, lumberjacks, silk weavers, and mechanics.³⁰ The Commission served as a training ground for a new generation of labor economists and social investigators who came into prominence during the 1920s and the New Deal.³¹

Sydenstricker also surveyed community welfare for the Commission and studied the public hearings on the building trades of New York City, collective bargaining in the anthracite coal industry in Pennsylvania, and the conditions of industrial workers.³² In his study of Pennsylvania coal miners, he argued that the causes of industrial unrest were the irregularity of employment, the fall of the real income of the miners, and “the manner of living, standard of home, the real value of the worker’s wages as expressed in commodities and enjoyment”; the wage earner lacked incentives to become a “stable citizen.”³³ The unrestrained growth of industrial centers had brought about poverty, physical degeneration, and moral deterioration, he argued.³⁴ Municipal governments, not social workers—with their “taint of charity”—should be responsible for community welfare, in his view.³⁵ His findings supported the Progressive reformers, who wanted the authorities to guarantee the rights of the poor to employment and insurance.³⁶

The Commission believed that the responsibility for the worsening conditions of workers and industrial violence lay with the big corporations: their enormous profits were not fairly shared with their workers or consumers. Instead of tracking real earnings over time, Sydenstricker’s report, *Conditions of Labor in the Principal Industries*, analyzed the extent to which a worker’s income would support a decent

standard of life: the average wage-earner’s family spent from three fourths to four fifths of its income on subsistence, leaving little or nothing for unemployment or old age.³⁷ According to Sydenstricker’s report, “Fully one-half of the wage-earners’ families in the United States do not have any income sufficient for adequate subsistence and health.”³⁸

While the Commission had not been formed to investigate the health of workers, in 1915, Warren asked Lauck to help the Service to determine the causes of sickness and mortality among wage earners in a typical industrial community. It rested on Warren’s idea that such problems had complementary “hygienic and economic” causes and, thus, the “economic and hygienic facts should be obtained at the same time and for the same conditions.”³⁹

A one-year study of “a fair type of the American industrial town” of 20 000 to 30 000 inhabitants, with 4000 workers in a single predominant industry and a complete range of wages, would be ideal, argued Warren: the community should be as average as possible,⁴⁰ and the study should focus on a cross-section of national conditions (although African Americans were excluded).⁴¹ The main value of that study, never made before, would be to show “the real” relation of economic and hygienic conditions in a typical industrial community.⁴²

Lauck fully agreed with the “determination in a scientific manner of the relation of sickness and mortality to poverty and all that the term poverty means.” In his opinion, although Rowntree’s study of poverty in York was groundbreaking, it was out of date and only applied to England; because “there is nothing in America of its kind,” it would be “the first survey in the field.”⁴³

Warren claimed that Sydenstricker's study for the Commission on the cost of living would have a direct bearing on the question of health, a study that "has never been made before;" because the whole subject "so closely involves economic facts and questions," Sydenstricker was hired by the Service.⁴⁴

THE SERVICE'S EARLIEST SICKNESS SURVEY

With Sydenstricker on board, Warren carried out his investigation of the causes of sickness and mortality among the wage earners of a typical industrial community. This plan drew on the Service's studies of pellagra in South Carolina, starting in 1914, under physician Joseph Goldberger. Goldberger sought to determine whether there was a causal relationship between diet and the disease with the use of clinical and experimental methods.⁴⁵ Warren and Sydenstricker planned to complement Goldberger's investigations with a study of the economic and sanitary factors responsible for pellagra, not just the dietary ones. For the 1915–1916 study, they chose seven cotton mill villages in northwestern South Carolina, with 500 to 800 inhabitants each, where pellagra was prevalent. A house-to-house canvass was supervised by physician George A. Wheeler, and the assembling of data on the families' diet and income by Sydenstricker. This field work (which continued until 1919 at least) also investigated the links between income and illness in general, not just with pellagra.

It was the first study of the close correlation of "hygienic" and "economic" conditions in an industrial community⁴⁶ and drew on the methods of Booth and Rowntree: the first-hand collection,

by a trained investigator, of data about individuals, families, and households on a local level, and an attempt to quantify the phenomena and pressure public policy to make social reforms.⁴⁷ Lauck believed that the Service's sickness survey would provide a "solid foundation for more intelligent and constructive remedial measures,"⁴⁸ while Warren and Sydenstricker urged the federal government to implement a system of sickness insurance, arguing that a study of the causes of sickness among wage earners would throw light on the responsibilities of employers, employees, and the public for these conditions.⁴⁹

Instead of the "indefinite terms" like "poor," "fair," and "well-to-do" commonly used to classify the population's living conditions, Sydenstricker and his associates used quantitative methods and, "for the purpose of accuracy and convenience," an index of living conditions was expressed in a numerical form.⁵⁰ They took into account the income and food requirements of the family when classifying the sample population, about 4160 people in about 750 households.⁵¹ The resulting groups—half-month income of less than \$6.00, \$6.00 to \$7.99, \$8.00 to \$9.99, \$10.00 to \$13.99, and \$14.00 or more—showed a crude case rate of pellagra per 1000 of 42.7, 26.0, 12.8, 4.1, and 3.4 for the year 1916, respectively. According to the authors, "It is the first reported study in which the degree of the long-recognized association between poverty and pellagra incidence is measured in a definite, purely objective manner."⁵²

Similarly, for the analysis of the possible association between family income and sickness in general, the same investigators used the term "disabling illness" as a proxy for sickness to make

their results as comparable as possible with the contemporary sickness surveys by insurance companies.⁵³ "Sick" persons were those who were "unable to work" on account of sickness or accident, including those "up and about but unable to work" and those confined to bed.⁵⁴ The community sickness surveys in question had been made by the Metropolitan Life Insurance Company in 1915. The definition of sickness as disability and the instructions to the field agents of the Service's survey were exactly those of the insurance company.⁵⁵ The unemployment surveys the insurance company made in conjunction with the US Bureau of Labor Statistics paved the way for the company's sickness surveys.⁵⁶

METROPOLITAN LIFE INSURANCE COMPANY

Unemployment had been one of the major concerns of the investigations of social unrest, not only of the Commission on Industrial Relations but also the Bureau of Labor Statistics and local organizations. The Bureau regarded unemployment as the greatest evil of the capitalist system. The Commission presented a novel official view of unemployment as a collective tragedy that society was responsible for, instead of voluntary organizations or the private sector.⁵⁷ Meanwhile, the Bureau, the government of New York City, and the Metropolitan Life Insurance Company set out to measure unemployment, first in New York City, then nationwide.⁵⁸

In January 1915, the agents of the Metropolitan, in conjunction with the Mayoralty of New York, visited the families of the policyholders, 155 960 in all, and concluded that 18% of all potential wage earners were unemployed. Having found the same percentages based

on the 1910 census, the Bureau enlisted the insurance company to make studies in other cities. All persons who had any employment, regular or irregular, full or part time, were recorded as employed.⁵⁹ The Bureau, in cooperation with the insurance company, surveyed 16 cities in the East and Midwest and 12 cities in the Rocky Mountains and the Pacific Coast to find out whether unemployment was because of “no work to be found,” “sickness or disability,” “strikes or lock-outs,” or “other causes.”⁶⁰

Having found that 11% of unemployment was associated with sickness or workplace accidents, the Metropolitan set out to measure the “amount of illness prevailing in American communities.”⁶¹ It chose the city of Rochester, New York, for its initial “sickness census” in September 1915. In all, 7638 families, made up of 34 490 persons, were surveyed, almost 10 times the number of families canvassed in the Service’s earliest sickness survey in South Carolina. The census divided sickness into cases in which the illness did or did not result in incapacity (to work). The instructions to the agents specify, for those who were found to be sick, that—

The sick should include: (a) Those persons who are up and about, but are unable to work because of sickness or accident; (b) Those who are confined to bed at home because of disease or accident; (c) Those who are receiving treatment in hospitals or similar institutions.

The question “how long sick to date” should be answered in days, weeks or months.⁶²

—the same method of the Service’s earliest community survey.

Disease understood as an inability to work surpassed disease understood as a specific medical diagnosis. Using the former, the company calculated the effects of sickness in the general population of Rochester, in terms of days lost from work and, thus, wages lost in a year. Illness in a community—or the prevalence of sickness—could thus be translated into economic losses—and perhaps into calculating premiums.⁶³

The company noted that “a sick benefit society experiencing a light sickness rate may come to grief if the sicknesses involve protracted payments of benefit out of all relation to the premiums charge.” So, classifying sickness rates by sex, age, and other measures was perhaps less important than

to know that, of those cases of sickness, 56% will probably last more than one month, 39% more than 3 months, 25% more than one year and 14% more than 3 years . . . persons interested in the social consequences of sickness are more likely to inquire into the durations of sicknesses than, primarily, into the sickness rate.⁶⁴

This definition of sickness dates back to the company’s free Visiting Nurse Service created in 1909. Lee K. Frankel, then head of the company’s welfare and health program, believed that the nurse service benefitted both the policyholder, usually a worker, and the company: the policyholders who were taught the importance of personal hygiene would be less likely to suffer from illness and premature death, and the company would likewise profit from a lower mortality rate.⁶⁵ But the company’s need to show that spending a policyholder’s money on health work would reduce the cost of insurance, with findings backed by data,⁶⁶ led it to

compile information about the policyholders, in the form of a “new case and history slip” the nurses had to fill in at the bedside, which included such questions as “how long was the patient ill at the time of the first visit? Up and about? In bed?”⁶⁷ This information was grouped into the general category of “inability to work” in the sickness surveys done in Rochester in 1915, which in turn were taken up by the Public Health Service’s sickness survey in the cotton mill villages in South Carolina. With this definition and metrics, the Service found in 1915 to 1916 that the rates of disabling illness (inability to work) per 1000 persons in the families of the cotton mill villages with an income of less than \$6.00, \$6.00 to \$7.99, \$8.00 to \$9.99, and \$10.00 and more were 70.1, 48.2, 34.4, and 18.5, respectively.⁶⁸

THE LATER SICKNESS SURVEYS

The socioeconomic classification of the families used by Sydenstricker in the South Carolina sickness survey, and the definition of illness proposed by Metropolitan Life Insurance Company, were at the core of several sickness surveys of the following decades. In the Service’s investigation of the mortality rates from influenza during the pandemic of 1918 and the sickness survey of Hagerstown, Maryland (1921–1924), of the incidence of several diseases among 1600 families (7200 persons), information on the economic status of the respondents was recorded but it was left to the enumerators to classify the families with the use of “indefinite”—not quantitative—categories: “well-to-do, comfortable, moderate, poor, very poor.”⁶⁹ The 1928 and 1931 surveys of 8758 families in 17 states by

the Committee on the Cost of Medical Care, with the participation of Sydenstricker and the insurance company, used the company's definition of illness: "any disorder which wholly or partially disables an individual for one or more days. . . ."70; and the number of days in bed, days lost from school or work, and total duration of each illness were also recorded.71

In the studies done during the Depression by the Milbank Memorial Fund and the Service with the Metropolitan Life Insurance Company under the auspices of the Health Organization of the League of Nations, quantitative data of 12 000 families in 10 localities, including eight large cities, were compiled. The households were classified according to their economic status as comfortable, moderate, or poor to determine and quantify whether a fall in income had affected their health.72 A "disabling illness" was defined as a situation in which "the person was unable to work, attend school or undertake other activities for 1 or more days during the 3 months period of the study."73 This methodology was based on the assumption that mortality statistics were an insufficient way to measure a nation's health. The conservative members of the League of Nations Health Organization, who eventually won the upper hand, had claimed that the decline in mortality rates during the 1930s meant that the economic crisis had not affected public health, while Sydenstricker, now working for the Milbank Memorial Fund, and George St John Perrott, from the Service, showed the opposite: that morbidity (disabling illness) was directly linked with socioeconomic status (income).74

Historian George Weisz has shown that defining disabling illness as inability to work and perform other activities

was crucial for the focus on chronic disease of the biggest morbidity survey up to then: the National Health Survey (1935–1936), whose aim was to support health care reform.75 Sydenstricker was the driving force of this survey as well. Weisz argued that its focus revealed the pervasiveness of chronic disease and how the poor, who had less access to health care, suffered more from diseases than other classes.76

CONCLUSIONS

The earliest socioeconomic sickness survey in an industrial community undertaken by the US Public Health Service in 1915, a result of the US social survey movement, marked the beginning of a methodology of research that stretched to the National Health Survey of the 1930s. Other agencies, like the US Children's Bureau and the National Tuberculosis Association, also used information on a family's economic status or occupation in their mortality investigations. With the decline of the Progressive movement, institutional and methodological changes, and the outbreak of World War II, the socioeconomic measurements used for vital statistics were virtually abandoned.77 However, the 1935–1936 National Health Survey's definition of illness in terms of days of disability played a significant role in the United States' epidemiological and political approach to chronic diseases, while the National Health Survey became a major source of information for the arguments that eventually led to the establishment of Medicaid.78 This definition originated with the Metropolitan Life Insurance Company's sickness surveys and was adopted by the surveys that the Public Health Service began in the 1910s.

Both organizations participated in the Committee on the Cost of Medical Care and in the studies made during the Depression. The extent to which the Service's concerns about the health of the wage earners coincided with the interests the company is a subject that requires further research. *AJPH*

ABOUT THE AUTHOR

Mónica García is with the School of Human Sciences, Universidad del Rosario, Bogotá, Colombia.

CORRESPONDENCE

Correspondence should be sent to Mónica García, School of Human Sciences, Universidad del Rosario, Cra. 6 # 12C-13, Off. 505, Bogotá 111711, Colombia (e-mail: claudia.garcia@urosario.edu.co). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONFLICTS OF INTEREST

The author reports no conflict of interest.

ENDNOTES

1. Nancy Krieger and Elizabeth Fee, "Measuring Social Inequalities in Health in the United States: A Historical Review, 1900–1950," *International Journal of Health Services* 26, no. 3 (1996): 391–418, <https://doi.org/10.2190/B3AH-Q5KE-VBGF-NC74>.
2. Stephen J. Mooney, Justin Knox, and Alfredo Morabia, "The Thompson-McFadden Commission and Joseph Goldberger: Contrasting 2 Historical Investigations of Pellagra in Cotton Mill Villages in South Carolina," *American Journal of Epidemiology* 180, no. 3 (2014): 235–244, <https://doi.org/10.1093/aje/kwt010>.

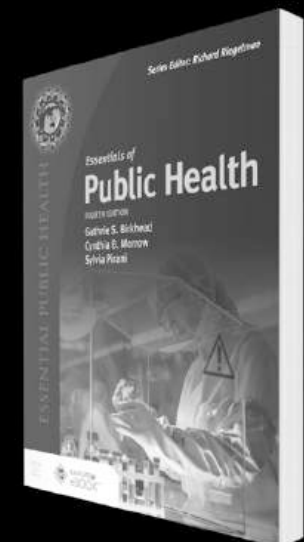
- 1093/aje/kwu134; John M. Eyler, "Health Statistics in Historical Perspective," in *Health Statistics: Shaping Policy and Practice to Improve the Population's Health*, eds. Daniel J. Friedman, Edward L. Hunter, and R. Gibson Parrish (New York, NY: Oxford University Press, 2005), 25–52; Paul Weindling, "Interwar Morbidity Surveys: Communities as Health Experiments," in *Facing Illness in Troubled Times*, eds. Iris Borowy and Wolf D. Gruener (Frankfurt am Main, Germany: Peter Lang, 2005), 75–84.
- Martin Bulmer, Kevin Bales, and Kathryn Kish Sklar, "The Social Survey in Historical Perspective," in *The Social Survey in Historical Perspective, 1880–1940*, eds. Martin Bulmer, Kevin Bales, and Kathryn Kish Sklar (Cambridge, England: Cambridge University Press, 1991), 1–48, 20–23.
 - Bulmer et al., "The Social Survey in Historical Perspective," 48; Eileen Janes Yeo, "Social Surveys in the Eighteenth and Nineteenth Centuries," in *The Cambridge History of Science*, Volume 7, *The Modern Social Sciences*, eds. Theodore Porter and Dorothy Ross (Cambridge, England: Cambridge University Press, 2003), 83–99.
 - Bulmer et al., "The Social Survey in Historical Perspective," 3.
 - On social survey, see Maurine W. Greenwald and Margo Anderson, *Pittsburgh Surveyed. Social Science and Social Reform in the Early Twentieth Century* (Pittsburgh, PA: University of Pittsburgh Press, 1996); for the Settlement House movement in the United States, see, for example, Harry P. Kraus, *The Settlement House Movement in New York City, 1886–1914* (New York, NY: Arno Press, 1980).
 - Allen Eaton and Shelby M. Harrison, *A Bibliography of Social Surveys* (New York, NY: Russell Sage Foundation, 1930).
 - The primary sources upon which this research was based are documents available at the National Archives at College Park, MD, and Harvard Libraries. At the National Archives, I checked the papers of the Commission on Industrial Relations where Sydenstricker worked before going to the US Public Health Service. They are located in the Microfilm Publications, T4, A2. These documents have been used by economic historians such as Mary O. Furner but never in relation to the Service's surveys. I also checked the papers and correspondence of the pellagra studies directed by the Service, where surgeon Benjamin S. Warren (advisor of the Service to the Commission on Industrial Relations) and Edgar Sydenstricker started their socioeconomic surveys. They are available at RG Central File 1897–1923, Pellagra, NC-34-10, File 1648, Boxes 148–153. I also checked the correspondence of the Service in RG 90 Insurance-Health, which refers to the initiative to make community sickness surveys and how it was related to the social survey. Finally, I also checked the National Institute Health Survey Forms, Hagerstown, in RG 90 Entry: A1 21, Boxes 1–20. As for the social survey and the Metropolitan Life Insurance Company, I relied on the historiography about the company as well as on published primary sources located at the Countway Library of Medicine and the Widener Library, both at Harvard University.
 - Thomas Bender, *A Nation Among Nations. America's Place in World History* (New York, NY: Hill and Wang, 2006), particularly chapter 5, "The Industrial World and the Transformation of Liberalism," 246–295; Richard L. McCormick, "Public Life in Industrial America, 1877–1917," in Eric Foner, *The New American History* (Philadelphia, PA: Temple University Press, 1997), 107–132; Greenwald and Anderson, *Pittsburgh Surveyed*; Graham Adams Jr, *Age of Industrial Violence, 1910–1915* (New York, NY, and London, England: Columbia University Press, 1966); Mary O. Furner, *Advocacy and Objectivity: A Crisis in the Professionalization of American Social Science, 1865–1905* (Lexington, KY: University Press of Kentucky, 1975).
 - Chantal Julia and Alain-Jacques Valleron, "Louis-René Villermé (1782–1863), a Pioneer in Social Epidemiology: Re-analysis of His Data on Comparative Mortality in Paris in the Early 19th Century," *Journal of Epidemiology and Community Health* 65, 2011: 666–670, <https://doi.org/10.1136/jech.2009.087957>.
 - Simon Szreter, "The GRO and the Public Health Movement in Britain, 1837–1914," *Social History of Medicine* 4, no. 3 (1991): 435–463, <https://doi.org/10.1093/shm/4.3.435>; and Simon Szreter, "The Official Representation of Social Classes in Britain, the United States, and France: The Professional Model and 'Les Cadres,'" *Comparative Studies in Society and History* 35, no. 2 (1993): 285–317.
 - Krieger and Fee, "Measuring Social Inequalities in Health," 394–395.
 - Eyler, "Health Statistics in Historical Perspective," 25–52, 39.
 - Szreter, "The GRO and the Public Health Movement," 435–463; and Szreter, "The Official Representation of Social Classes," 285–317.
 - Krieger and Fee, "Measuring Social Inequalities in Health," 394–399.
 - Sydenstricker is listed as one of the graduate students of the University of Chicago in James Lawrence Laughlin, *Twenty-Five Years of the Department of Political Economy* (University of Chicago, Priv. Print., 1916), 16. See also Wilford I. King, "Edgar Sydenstricker," *Journal of the American Statistical Association* 31, no. 94, 1936: 411–415; Washington and Lee University, *The Calix 1902. Yearbook*, Vol VII (Louisville, KY: The Prentice Press, 1902), 25. What we know about Edgar Sydenstricker derives from his work for several institutions. For his work for the Commission on Industrial Relations, see Mary O. Furner, "Knowing Capitalism: Public Investigation and the Labor Question in the Long Progressive Era," in *The State and Economic Knowledge. The American and British Experiences*, eds. Mary O. Furner and Barry Supple (Cambridge, England: Cambridge University Press, 1990), 275. About his work for the US Public Health Service, see Krieger and Fee, "Measuring Social Inequalities in Health," and Harry Marks, "Epidemiologists Explain Pellagra: Gender, Race and Political Economy in the Work of Edgar Sydenstricker," *Journal of the History of Medicine* 58, no. 1 (2003): 34–55, <https://doi.org/10.1093/jhmas/58.1.34>. We also know that he assisted the League of Nations Health Organization as a statistical adviser: see Weindling, "Interwar Morbidity Surveys," 77. He also helped in the design of the National Health Survey of 1935–1936: see George Weisz, "Epidemiology and Health Care Reform. The National Health Survey of 1935–1936," *American Journal of Public Health* 101, no. 3 (2011): 438–447, <https://doi.org/10.2105/AJPH.2010.196519>. Finally, we also know that while working for the Milbank Memorial Fund, he advocated for national health insurance before the Committee on the Cost of Medical Care and the Committee on Economic Security of the Roosevelt administration: see Jaap Kooijman, . . . *And the Pursuit of National Health. The Incremental Strategy Toward National Health Insurance in the United States of America* (Amsterdam, Netherlands: Rodopi, 1999).
 - Sydenstricker's work for the Immigration Commission was most likely his first experience of fieldwork on the condition of industrial workers and the use of quantitative methods and descriptive accounts in industrial surveys. The purpose of the Immigration Commission was to study the origins of the immigration of the 1880s, motivated by the higher wages paid for industrial labor and its effects upon the institutions and industries of the country—see *Reports of the Immigration Commission. Abstracts of Reports of the Immigration Commission: With Conclusions and Recommendations and Views of the Minority*, V1 (Washington, DC: Government Printing Office, 1911), 13. The Commission reflected the spirit of Progressivism, a movement to improve the country's socioeconomic conditions, relying on what their adherents considered to be scientific objectivity and a commitment to government intervention. The Progressives believed that investigations carried out by properly trained experts, not by well-meaning amateurs or charities, would equip policymakers with the means to solve social problems—see Robert F. Zeidel, *Immigrants, Progressives, and Exclusion Politics. The Dillingham Commission, 1900–1927* (DeKalb, IL: Northern Illinois University Press, 2004), 3. Two of the nine-member Commission were professors of political economy, and the support staff included the political economist William Jett Lauck, who was trained by James Laurence Laughlin and Thorstein Veblen at the Department of Political Economy of the University of Chicago—see Laughlin, *Twenty-Five Years of the Department of Political Economy*, 9. One of Lauck's assistants was Edgar Sydenstricker, then a graduate student of political economy at the University of Chicago—see Laughlin, *Twenty-Five Years of the Department of Political Economy*, 16; see also King, "Edgar Sydenstricker." The Commission obtained original information on about 3.2 million people. Their fieldwork was done in 24 states with 17 141 households and studied several industries, from sugar refining to coal mining. The investigations combined data "which were susceptible of tabulation and statistical presentation" about the individual employee, his household, and industrial plant, along with descriptive and historical information about 200 industrial plants and communities: see *Reports of the Immigration Commission*, 293–295.
 - These sectors were the "professional classes," that is, women of a privileged social position, professionals in the field of medicine, members of voluntary welfare associations, some journalists, and, at the end of the century, a few academics responsible for "social surveys." See Bulmer et al., "The Social Survey in Historical Perspective," 2.
 - Benjamin Seebohm Rowntree, *Poverty. A Study of Town Life* (London, England: Macmillan and Co Limited, 1901), vi.
 - Charles Booth, *Life and Labour of the People in London. Volume I. East, Central and South London* (London, England: Macmillan and Co, 1892), 6.
 - Bulmer et al., "The Social Survey in Historical Perspective," 20–21.

22. Jean M. Converse, *Survey Research in the United States. Roots and Emergence, 1890–1960* (New Brunswick, NJ: Transactions Publishers, 2009), 22–23; Martin Bulmer, “The Decline of the Social Survey Movement and the Rise of American Empirical Sociology,” in Bulmer et al., *The Social Survey in Historical Perspective*, 991–315, 293.
23. Steven R. Cohen, “The Pittsburgh Survey and the Social Survey Movement: A Sociological Road Not Taken,” in Bulmer et al., *The Social Survey in Historical Perspective*, 245–268, 247–249.
24. Adams, *Age of Industrial Violence*, 25–30.
25. *First Annual Report of the Commission on Industrial Relations* (Chicago, IL: Barnard & Mills Print, 1914), 6.
26. Lauck graduated from Washington and Lee University and obtained a degree in political economy at the University of Chicago. See Laughlin, *Twenty-Five Years of the Department of Political Economy*, 9; Washington and Lee University, “University, Washington, D.C.,” *The Ring-Turn Phi* 11, no. 17 (1908): 2; Washington and Lee University, “Hart, Schaffner & Marx Prizes,” *The Ring-Turn Phi* 12, no. 11 (1909): 4.
27. Washington and Lee University, “Dr G. Campbell Tells of Industrial Commission. Vitality Important Work Being Done by W. & L. Men,” *The Ring-Turn Phi* 17, no. 18 (1914): 3.
28. Benjamin S. Warren, “Outline of Plan for the Study of Sickness Insurance,” April 23, 1915, in National Archives, RG 90 Insurance-Health, 6. Elizabeth Fee’s and Theodore M. Brown’s assertion that the available biographical information about Benjamin Warren is very meager still holds true: see Elizabeth Fee and Theodore Brown, “Edgar Sydenstricker and Benjamin Warren,” *American Journal of Public Health* 89, no. 11 (1999): 1643–1644, 1643. Warren’s participation in some of the Service’s projects, such as the pellagra studies, is mentioned in Ralph Chester Williams, *The United States Public Health Service, 1789–1950* (Washington, DC: US Public Health Service, 1951), 152, 423. To my knowledge, this article is the first to discuss his work as the sanitary adviser of the Commission on Industrial Relations and his role in implementing the sickness survey among wage earners.
29. *First Annual Report of the Commission*, 8.
30. Adams, *Age of Industrial Violence*, 206; *Reports of the Immigration Commission*, 2.
31. Mary O. Furner, “Knowing Capitalism: Public Investigation and the Labor Question in the Long Progressive Era,” in *The State and Economic Knowledge. The American and British Experiences*, eds. Mary O. Furner and Barry Supple (Cambridge, England: Cambridge University Press, 1990), 241–286, 275.
32. Edgar Sydenstricker, “Topical Analysis of Testimony at Public Hearings on the Building Trades of New York City, N.Y. May 25–29, 1914,” in Microfilm publications, T4, Reel 14, A2 Cab 41/4, 50 pages; Edgar Sydenstricker, “Report on the System of Collective Bargaining in the Anthracite Coal Industry,” in National Archives, Microfilm Publications, T4, A2 Cab 41/4, Reel 8. See also Edgar Sydenstricker, “Condition of Labor in the Principal Industries,” s/f, in National Archives, Microfilm Publications, T4, A2 Cab 41/4, Reel 9. This served as the basis for W. Jett Lauck and Edgar Sydenstricker, *Conditions of Labor in American Industries* (New York, NY, and London, England: Funk & Wagnalls Company, 1917).
33. Edgar Sydenstricker, “Appendix M. Industrial Communities,” in National Archives, Microfilm Publications, T4, A2 Cab 41/4, Reel 11, 1–11, 2.
34. Edgar Sydenstricker, “Welfare Activities of Communities. Report to Sept. 1, 1914,” in National Archives, Microfilm Publications, T4, A2 Cab 41/4, Reel 1, 1.
35. Sydenstricker, “Welfare Activities,” 2–3; H. G. Lee, “Digest of Mr. Sydenstricker’s Preliminary Report on Welfare Activities of Communities,” February 18, 1915, in National Archives, Microfilm Publications, T4, A2 Cab 41/4, Reel 11.
36. Social progressives, under the American Association for Labor Legislation, played an important role in establishing workmen’s compensation and the campaign for health insurance. See Paul Star, *The Social Transformation of American Medicine* (New York, NY: Basic Books, 1982), 243–249; Kooijman, . . . *And the Pursuit of National Health*, 25–30.
37. Furner, “Knowing Capitalism,” 278.
38. Sydenstricker, “Condition of Labor in the Principal Industries,” 1.
39. Benjamin S. Warren, “Industrial Hygiene. Preliminary Outline of Plan for the Study of a Typical Industrial Community,” January 19, 1915, in National Archives RG 90 Insurance-Health, 1.
40. To ensure a fair sampling—or the “representative method” as it was called—the American social surveys based their selection of communities on the idea of a “typical” village or “average” community, as Charles Booth did when he used the School Board Visitors’ data on the poor in London, England, and the Service when it surveyed the cotton-mill villages in South Carolina. On the representative method in US social survey, see Alain Desrosières, “The Part in Relation to the Whole: How to Generalize? The Prehistory of Representative Sampling,” in Bulmer et al., eds., *The Social Survey in Historical Perspective*, 217–244, 220.
41. Warren excluded African Americans because he wished to model his survey on an “average” community; he argued that too great a diversity of races among the wage earners should be avoided because of variations in standards of living caused by “racial customs.” He was also concerned about the impact of his conclusions: he wanted to avoid any “appearance of muckraking,” on the one hand, or present better-than-ordinary conditions, on the other. For him, “complicating” factors should be avoided; hence, as far as possible, the study had to be based on typical conditions. See Warren, “Industrial Hygiene,” 2, 8. Nevertheless, historian Harry M. Marks has shown that Sydenstricker’s focus on economic underdevelopment in the Service’s pellagra studies led him to ignore that African Americans were the main victims of the disease. This indifference, he argues, also reflected a broader concept of race at that time. For Northerners, race meant “immigrants.” When settlement workers and the social economists of Sydenstricker’s cohort spoke of “races,” they were generally referring to recent immigrants. See Marks, “Epidemiologists Explain Pellagra,” 53–55.
42. Warren, “Industrial Hygiene,” 2, 8.
43. Letter by Jett Lauck to Benjamin S. Warren, Chicago IL, February 17, 1915, National Archives RG 90 Insurance, Health, 1–2.
44. Benjamin S. Warren, “Outline of Plan for the Study of Sickness Insurance,” April 23, 1915, in National Archives, RG 90 Insurance-Health, 6 pages, quotations in pages 1, 2, and 6 respectively. Sydenstricker was the first statistician of the service. See letter from Rupert Blue to the Secretary of the Treasury, July 19, 1915, in National Archives, RG 90 Entry NC 34-10 Box 292 File Number 3339 to 3343.
45. Mooney et al., “The Thompson-McFadden Commission”; Marks, “Epidemiologists Explain Pellagra,” 34–35.
46. Letter by GA Wheeler, P.A. Surgeon, Medical Officer, in Immediate Charge, to Surgeon General, USPH Service, Field Investigations of Pellagra, Spartanburg, SC, December 10, 1919, National Archives, RG 90 Central File 1897-1923, Pellagra, NC-34 10, File 1648, Box 148 to 153, Box 151; Letter by Joseph Goldberger, Surgeon in Charge of Pellagra Investigations, to Surgeon General, USPH Service, Washington, July 2, 1920, National Archives, RG 90 Central File 1897–1923, Pellagra, NC-34 10, File 1648, Box 148 to 153, Box 151.
47. Sydenstricker and his collaborators referred to Seeborn Rowntree’s work on poverty in York, England, as a “classic” when discussing the measurement of the economic status of families: Edgar Sydenstricker and Wilford I. King, “A Method of Classifying Families According to Incomes in Studies of Disease Prevalence,” *Public Health Reports* 35 no. 48 (1929): 2829–2846, 2833; and Edgar Sydenstricker and Wilford I. King, “The Measurement of the Relative Economic Status of Families,” *Quarterly Publications of the American Statistical Association* 17, no. 135 (1921): 842–857, 844. They even used the information gathered by the Service’s earliest sickness survey in South Carolina cotton-mill villages to ascertain the variations in economic status at different stages of family life, as Rowntree had done in his book and as a direct application of Rowntree’s idea: Edgar Sydenstricker, Wilford I. King, and Dorothy Wiehl, “The Income Cycle in the Life of the Wage-Earner,” *Public Health Reports* 39, no. 34 (1924): 2133–2140.
48. Letter by Jett Lauck to Benjamin S. Warren, Chicago IL, February 17, 1915, National Archives, RG 90 Insurance, Health, 1.
49. Benjamin S. Warren and Edgar Sydenstricker, “Health Insurance. Its Relation to the Public Health,” *Public Health Bulletin* 76, 2nd ed (Washington, DC: Government Printing Office, 1916), 38.
50. Edgar Sydenstricker and Wilford I. King, “A Method of Classifying Families According to Incomes in Studies of Disease Prevalence,” *Public Health Reports* 35, no. 48 (1920): 2829–2846, 2830.
51. The size, ages, and genders of the families were taken into account by employing a common denominator for individuals, derived from the Atwater scale of food requirements. Thus, Sydenstricker and his colleagues set 1 as the requirement of a male adult, while the requirements of children and females were set at between 0.2 and 0.8. The 747 households were then classified according to a half-month family income per adult male unit with data canvassed between April and June, 1916. Joseph Goldberger, G.A. Wheeler, and Edgar Sydenstricker, “A Study of the Relation of Family Income and Other Economic Factors to Pellagra Incidence in Seven Cotton-Mill Villages of South Carolina in

- 1916," *Public Health Reports* 35, no. 46 (1920): 2673–2714, 2678–2684.
52. Goldberger et al., "A Study of the Relation of Family Income," 2687. Quote on 2711.
 53. Edgar Sydenstricker, G. A. Wheeler, and Joseph Goldberger, "Disabling Sickness Among the Population of Seven Cotton Mill Villages of South Carolina in Relation to Family Income," *Public Health Reports* 33, no. 47 (1918): 2038–2051.
 54. Sydenstricker et al., "Disabling Sickness," 2038–2039, 2043–2046.
 55. John Eyler also mentions this connection in his chapter "Health Statistics in Historical Perspective," 39–40.
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“American Indian” as a Racial Category in Public Health: Implications for Communities and Practice

Danielle R. Gartner, PhD, Rachel E. Wilbur, MPH, and Meredith L. McCoy, PhD, MEd

When public health considers the health and disease status of Indigenous people, it often does so using a racial lens. In recent decades, public health researchers have begun to acknowledge that commonly employed racial categories represent history, power dynamics, embodiment, and legacies of discrimination and racism, rather than innate biology. Even so, public health has not yet fully embraced an understanding of other components of identity formation for Indigenous people, including political status within Native nations.

In this article, we discuss why the continued racial conceptualization of Indigeneity in US public health is inadequate. We begin by providing a brief account of racialization as a tool of colonization, of failure to recognize and acknowledge Indigenous sovereignty, and of common public health practices of Indigenous data collection and interpretation.

We then articulate the stakes of racialized health data for Native communities. We end by offering alternative approaches, many drawn from scholarship from Indigenous researchers. (*Am J Public Health*. 2021;111(11):1969–1975. <https://doi.org/10.2105/AJPH.2021.306465>)

In what is currently the United States, it is common to encounter statements like “American Indian/Alaska Native adults are nearly three times more likely than non-Hispanic white adults to be diagnosed with diabetes,”¹ and “the incidence of laboratory-confirmed COVID-19 cases among AI/AN persons was 3.5 times that among white persons.”^{2(p1169)} In addition to communicating the disproportionate disease burden borne by American Indians, Alaska Natives, Native Hawaiians, and Pacific Islanders in the United States, these aggregated, racially bound statistics are difficult to accurately interpret and carry four significant implications, which we outline in the following

paragraphs. We primarily focus this article on the experiences of American Indian people whose traditional homelands are located in what is currently understood as the lower 48 states of the United States. Although many of these experiences will also resonate with Native Hawaiians, Pacific Islanders, and Alaska Natives, there are also critically distinct contexts that shape issues of citizenship and governance differently for their communities.

In this article, we suggest that without careful contextualization and consideration of the historical background of racialization in the United States, common practices that categorize Indigenous people by race imply a biological

or genetic origin, rather than a structural cause, for health differences between Native and non-Native communities. Second, by ignoring the unique and complex sociopolitical and historical contributors to health within each tribal community, such approaches mistakenly represent diverse Native communities as a monolith. Third, when researchers perceive Native peoples as a homogenized group, they often fail to assess the importance of tribal sovereignty and locally specific metrics of belonging within Native nations. Lastly, such approaches obscure how colonial policies have disproportionately and negatively affected health and well-being,

placing the blame on individuals instead of the compounded impacts of a systematic lack of access to care, environmental racism, economic deprivation, educational violence, and discriminatory policies.

We begin with background information about identity, racialization, and common practices for collecting and using data about Native people from public health research. After detailing what is at stake for Native communities when collection and interpretation of Native-specific health data rely solely on a racial lens, we end by offering alternative approaches.

Throughout this article, we use “Native” and “Indigenous” interchangeably, a practice based on our experiences as Indigenous women and as researchers. The term “American Indian” is inscribed in federal law, and we use American Indian or the category “American Indian/Alaska Native” (or AI/AN) where such categorization aligns with large existing data sets or with federal law.

BACKGROUND

Before the onset of colonization, there were no “Indians” or “Natives”; rather, there were Anishinaabeg, Dee-ni’, Chahta, Diné, Samoans, and hundreds of other distinct societies. Our shared experiences of colonization link us together, even as our distinctiveness as individual nations persists. Our identities today continue to be intimately shaped by our relationships to each other, to our homelands and waters, to our plant and animal relatives, to our languages, and to our teachings. We do not, and have never, needed external governments or researchers to tell us who we are. Rather than biology, our identities reflect our shared obligations

to one another and these ongoing connections.

For many Native people, our status as citizens or descendants of specific Native nations plays a key part in how we perceive our identities. This forms the basis for our political status, as citizens of both the United States and our respective Native nations. Native nations are sovereign governments with the inherent right to determine criteria for their citizenry, a right that the Supreme Court of the United States (in *Santa Clara Pueblo v Martinez*) and the United Nations (in its Declaration on the Rights of Indigenous People) have each recognized as among the core obligations and rights of any sovereign government. Each Native nation makes decisions about its citizenry based on its unique historical, sociopolitical, and environmental contexts.³

Today, Native people live in urban, suburban, and rural settings across all states and territories. As researchers, it is imperative that we listen to, honor, and abide by the ways in which people identify themselves, including attending to the vast diversity within and between Native communities. Doing so must take priority over convenience in how we think about, organize, analyze, and present our data. Doing otherwise is not only unethical, but can also result in poor, inaccurate results that collapse critical differences between communities.⁴

FEDERAL ATTEMPTS TO RACIALIZE AND HOMOGENIZE

The United States, a nation with a long history of differentially defining racial groups using biologically based notions of race to the direct benefit of White Americans, has insisted on a false

homogenized and racialized narrative about Indigeneity.³ Blood, for example, has been used as a structuring category to both categorize and erase Native people. This has been most visible in the policy of blood quantum, a government-created metric under which agents of the government would identify and record Native people by their perceived percentage of “Indian blood,” often based on phenotypic characteristics. Use of blood quantum was the beginning of a system of administrative genocide whose purpose was to eliminate Indigenous peoples and, by extension, open Indigenous lands for settler use.

This homogenization has long been a tool for settler acquisition of Indigenous lands; during the treaty-making period, policymakers used blood as a metaphor for identity as they defined legal and social concepts of race.⁵ The United States used this racialization in its attempts to render Native peoples interchangeable for the purposes of control, as federal Indian Affairs Commissioner J. D. C. Atkins wrote in 1888: “The object of greatest solicitude should be . . . to blot out the boundary lines which divide [Native people] into distinct nations, and fuse them into one homogeneous mass.”^{6(p9)} Rather than shared biological traits, it is this shared experience of colonial harm, as well as shared strategies for resistance to it, that connect Native people.

PUBLIC HEALTH METRICS AND INDIGENEITY

The social and historical context of each Native nation shapes the distribution of health and disease within its population today, because of centuries of colonial policies aimed at Indigenous

elimination.^{7,8} Examples of these include, among others, removing Native people from their homelands, which disrupted medicinal, nutritional, and ethnobotanical practices; crowding Native students into federal boarding schools, deeply unhygienic spaces where Native youths faced both assimilation and abuse as well as disconnection from their families, communities, and traditional knowledge; sterilizing Native women in Indian Health Service facilities to reduce the size of subsequent generations; mining, nuclear testing, and nuclear waste disposal on the homelands of American Indians, Alaska Natives, Native Hawaiians, and Pacific Islanders, resulting in high rates of environmental exposures and disease; and superimposing government-delivered commodity food rations while restricting access to traditional Native foodways, contributing to epidemics of diabetes and metabolic disease in Indian Country.

Existing health disparities stem from these and other forms of structural harm rather than biological deficiency. And yet, the persistent rhetoric of race within public health research obscures the impact of colonial violence on health outcomes. To make matters more complicated, there is no consensus for how to collect, organize, and analyze data from Indigenous research participants. Large data sets often use either self-reported survey questions, which may allow respondents to include their tribal nation in addition to the more generic “American Indian/Alaska Native,” or close-ended questions that ask respondents to identify themselves within a given racial schema (e.g., White, Black, other). Both options are vulnerable to overcounting (when people who are not recognized by a Native community mark themselves as

Native) and undercounting (when multi-racial people are not given the option to select multiple racial identities). As an example, the 2020 US Census, which helps determine resource allocation based in part on race, used the following close-ended self-reported categories for race: White, Black or African American, American Indian or Alaska Native, Asian, Native Hawaiian and Pacific Islander, and some other race, with the additional option to write in tribal affiliation. Other studies or surveys measure race as determined by an outsider, often a researcher or medical professional, with or without the input of the survey participant. This can result in discrepancies between an individual's self-identification and their official record, as stereotypes of what Native people are “supposed to look like” dominate the national imagination.

WHAT IS AT STAKE

As the categories “American Indian/Alaska Native” and “Native Hawaiian/Pacific Islander” have become codified in many areas of social inquiry in the United States, so too have they passed into usage in public health research, leading many researchers to erroneously see racial articulations of Native peoples as legitimate. To better express the problems associated with applying these racial categorizations, we name what is at stake for Native people and communities in the United States with this continued use.

Public health's use of racial categorization ignores the treaty-based nation-to-nation relationship between the United States and Native nations. Doing so also ignores the legal implications of sovereignty and self-governance, particularly around defining citizenship.⁵ Any attention to Indigenous health outcomes

must therefore reflect the ways in which Native nations define their citizenry, rather than masking each under a ubiquitous, homogenizing category.

The continued application of this categorization also ignores individuals' articulations of themselves. Use of the racialized category is a values-based decision, one that sends the message that the racial designation of Indigenous people matters more than other ways of defining Indigeneity.^{9,10} Privileging race means that other articulations of belonging, based on ceremony, citizenship, or kinship, are not seen as legitimate or are seen as less legitimate than racial criteria.⁹ Continued use of the racial Indigenous categorization lends credibility to the colonizer's view of who qualifies as “Indian” and clearly articulates who has the power to decide.^{10,11} Any attention to Indigenous health outcomes must therefore also reflect the ways in which Indigenous peoples define themselves.

Public health's use of the racial categorization system also supports the dispossession of Native communities of their land and resources,^{10,11} as it furthers the logic embedded in colonial policies that have long been “rooted in imperialism, White supremacy, and a desire for material gain.”^{12(p92)} Indigenous populations, when viewed as a racial category, are seen as small and shrinking in size because of inaccurate assumptions about the inevitability of Indigenous extinction, assumptions that emerge from both historical narratives steeped in Manifest Destiny and racist beliefs that multiracial Indigenous people are not “Indigenous enough.” Research practices that lump or altogether drop Indigenous peoples from data sets through practical considerations regarding sample size perpetuate

BOX 1— Potential Solutions for Commonly Encountered Issues in Indigenous Health Research: Data Aggregation

- Data aggregation severely limits the local applicability of research results but is unavoidable if no effort is made to obtain tribal affiliation of participants.
 - Research that is conducted by or in collaboration with Native nations, on health concerns directly identified by those nations, is more likely to focus on specific tribal communities instead of a homogenized “Native America.”
 - When feasible, data collection should allow for statements of tribal affiliation, including descendency.¹⁷ In these cases, use of “tribally affiliated” rather than “tribal citizenship” is preferred to account for unenrolled people who maintain tribal affiliation.¹⁸
 - Tribe-specific data collection greatly assists in the development of local programs and policies and aids in the disaggregation of aggregated estimates.¹⁹ However, when performing secondary data analyses, it may prove impossible to differentiate between tribal communities if the original survey failed to collect such data. In these cases, the inability to specify findings by tribal group should be listed as a limitation.
 - The inability to differentiate between tribal nations is a common limitation of secondary data analyses. Recognition of data limitations should motivate researchers to “do better” in the future, improving the quality of tribal data in public health over time by oversampling in population-wide studies or by conducting tribe-specific studies.¹⁶
- Data aggregation may raise ethical concerns if appropriate tribal approvals are not sought during the research process.
 - Tribal IRBs (institutional review boards) and RRBs (research review boards) exist to protect the interest of tribes, particularly given previous, egregious research-based mistreatment of Indigenous peoples. Therefore, it is essential to engage with tribal research boards and Tribal Councils from the tribal nation of each participant when conducting research with Native peoples.^{16,18,20}
 - Do not publicly report on tribe-specific information without consulting Tribal Councils²¹ and having data use agreements in place.¹⁸
- Data aggregation across tribes obscures important health disparities within and between tribes.²²
 - Each Native nation has unique community values, history, religion, language, epistemology, political affiliations, presence of economic drivers, age structure, health insurance coverage and access, disease prevalence, distributions of risk, and protective factors. Each of these affects the distribution of health and disease; policies and programs aimed at eliminating health disparities should be tailored to local contexts to achieve greatest effectiveness.
 - There are instances, however, in which studying the aggregate Native experience is called for: specifically, when assessing the impacts of structural, environmental, and interpersonal racism on Native nations collectively. In these cases, it is the racism of a system and the outcomes of that racism that are being evaluated, not the “race” of the people that system affects. Approval from tribes whose data are being aggregated is still required.²⁰
 - Other instances in which data aggregation may be appropriate include the following²⁰:
 - Similarities in community locality and characteristics, including population density (urban/rural) and availability of natural resources (coastal, inland).
 - Similarities in environmental contaminant exposure that may affect local food, water, and soil.
 - Similarities in ability to access health-related resources and services, including travel distances to resources and services, cultural sensitivity of care provision, spaces for physical activity, healthy food availability, and educational opportunities.
 - Overall affluence and economic development, including presence of local economic drivers and unemployment levels.

discourses of Indigenous disappearance and death by reflecting a false perception of a country without Native peoples where Native lands and resources are up for grabs by others.

Additionally, non-White racial categories are often conflated with innate biological or genetic deficiency, and thus any connection to disease is highlighted.¹³ As such, racialized

communities are seen as bearing a disproportionate burden of morbidity and mortality, whereas Whiteness is rarely connected with health risks.¹⁴ One result of incorrectly placing causation

BOX 2— Potential Solutions for Commonly Encountered Issues in Indigenous Health Research: Avoiding the “Other” Category

- Adequate data regarding disease burden is critical when justifying the need for prevention or intervention programs and health policies. Indigenous peoples may make up a small proportion of the overall participants in health studies, but altogether removing Indigenous peoples from results or collapsing Indigenous populations into an “other” category obscures important and needed health data.
- Studies should report and publish results for numerically small groups, regardless of statistical significance of estimates.¹⁷ For a discussion of similar dynamics in higher education research, see Shotton et al.²³
 - When sample sizes do not allow for statistical significance to be achieved, studies should avoid the urge to group numerically small groups, including Indigenous communities, into an “other” category. Instead, report results by group, regardless of sample size.¹⁷ Indeed, when practitioners attempt to use existing data to justify the need for programs or to develop prevention or intervention projects, estimates that categorize Indigenous communities as part of an “other” group are of little utility.¹⁷
 - Limited data, when limitations are clearly specified, are better than no data. Withholding Indigenous-specific estimates makes it impossible for future meta-analyses or pooling projects to make use of existing study results. Further, Indigenous communities have noted that although they participate in studies, they often do not see results that are relevant to them. This weakens relationships necessary for the research enterprise to be sustained. Additional useful information includes sample size (raw numbers), percentage of total population, and confidence intervals (with explanations of their meanings). Include issues that stem from sample size in the limitations section.
 - Small sample size should not be considered an obstacle to high-quality research, particularly since findings can yield useful information for local practitioners. Include the results of qualitative techniques, such as focus groups or talking circles, to provide important information otherwise obscured by small sample sizes.¹⁸

BOX 3— Potential Solutions for Commonly Encountered Issues in Indigenous Health Research: Indigenous Identification and Classification

1. Providing an opportunity to mark “all that apply” rather than a single racial or ethnic identity allows individuals free choice in self-identification.
 - Several governmental data collection efforts (e.g., the US Census) include options to self-identify and select multiple racial and ethnic identities.
 - Carefully consider when it is best to have participants identify as multiracial and how those individuals will be considered in analyses and reporting.¹⁷ Make your intentions clear and include your rationale in research communications.¹⁶
 - Some community advocates have suggested that Indigenous peoples check only 1 race to ensure that they are tabulated as Indigenous and that associated resource distributions would be more equitable.¹⁷ Always consult and seek guidance from Indigenous communities to better understand how they define themselves.²⁴
 - “Select all that apply” instructions may lead to highly specific results or groupings that are difficult to interpret (i.e., “American Indian and Black” vs “American Indian, Black, and Asian”); it is still important, however, to offer the opportunity to self-identify. Aggregation into broader categories can be done, with documentation.¹⁸
2. Racial misclassification, specifically undercounting of Indigenous populations, is a well-described issue in health research. Administrative data frequently misreport Native people as belonging to other racial or ethnic groups.
 - Under no circumstances should phenotypic or cultural characteristics (hair, skin color, clothing, language) be used to assign racial identity or tribal affiliation.²¹
 - Racial classification by others (i.e., not self-identified) is affected by social processes and stereotypes and is therefore not accurate.^{4,21}

for the burdens of disease on racial groups instead of structural factors is that it blames Native people (rather than settler colonialism) for health disparities.

Finally, condensing all Native people into a single category erases the unique experiences of each tribal community—experiences that have a significant impact on health. Conclusions and health interventions based on research conducted using such broad strokes are unlikely to be effective at either identifying the root causes of health disparities unique to each tribal community or identifying effective intervention strategies. In light of the immense diversity within Indian Country, Native people should not be consolidated under a single category for health research. Such

statistical categories obscure the distinctions between Native nations (each with its own histories, political and environmental contexts, and systems for governance, education, and health) and reify notions of Indigenous essentialism.

TOWARD SOLUTIONS

In this article, we have offered a brief window into identity and the ramifications of public health’s “colonial blindness” on the collection and analysis of health data for Indigenous peoples.¹⁵ Moving ahead, we urge a pivot in public health scholarship to an approach that identifies the true factors responsible for poor health outcomes in Native communities. These shifts may lead to more targeted, systems-focused

interventions and promote increased tribal self-determination in health research processes. For those interested in pursuing public health research with Indigenous communities, we offer a set of recommendations, which join our own suggestions with those previously published by colleagues.^{16,17} Indeed, 20 years ago, Burhansstipanov and Satter raised several of these issues and offered solutions in this very journal.¹⁷ We lament that many of these suggestions have gone unheeded. Boxes 1–4 synthesize previously proposed solutions as well as our own suggestions. Note that the boxes do not provide an exhaustive list of all activities that researchers might do to be good stewards of the research process, nor do they reflect every recommendation from previous studies.

BOX 4— Potential Solutions for Commonly Encountered Issues in Indigenous Health Research: Avoiding Exploitative and Extractive Research Practices

There is a long history of researchers enacting exploitative and extractive practices in Indigenous communities, and studies have too often presented Indigenous peoples as deficient.

- The foundation of Indigenous data sovereignty is that Indigenous peoples have the right to determine how, when, where, and for what purposes data about them are collected, accessed, analyzed, interpreted, managed, and disseminated. Because of this, researchers must pay attention to local data sovereignty practices and consult data sovereignty networks.^{25–28}
- Data needs vary across Indigenous peoples. Data are needed that meet the needs and desires of specific communities. When Indigenous data needs are met, and data reflect the priorities, values, and diversity of Indigenous peoples, then these data can disrupt deficit narratives, present the realities of Indigenous people’s lives, and support Native nation (re)building.^{26,29}
- Research findings must benefit the participating community.²⁷ Findings should be shared with the participating community before being disseminated to the broader scientific community.^{20,25}

However, we believe the included recommendations offer a good place to start in supporting Indigenous public health in a way that upholds tribal nations' sovereignty with regard to their people, lands, and data. Such recommendations are in alignment with Articles 4 and 7 of the declaration of the Indigenous and Tribal Peoples Convention of the International Labor Organization, which hold that Indigenous peoples have the right to set priorities and parameters for projects that affect their lives, and that all initiatives must protect against harm to Indigenous bodies, lands, properties, cultures, and institutions.

CONCLUSION

Public health researchers and practitioners must listen to and honor the ways in which Indigenous peoples identify themselves. Doing so must take priority over convenience in how we think about, organize, analyze, and present our data. Given the history of unethical and inappropriate health research in Indian Country and the long-standing practice of non-Native researchers imposing their own definitions for and expectations about Native communities, it is essential to listen to and follow the research priorities and guidelines of Native nations and Native-run organizations. Native-run organizations have issued guidance for these processes, including the National Congress of American Indians' *Walk Softly* report,²⁷ several briefs by the Urban Indian Health Institute,³⁰ and the United States Indigenous Data Sovereignty Network.³¹ Relatedly, a recent National Institutes of Health–commissioned document, *American Indian and Alaska Native Research in the Health Sciences: Critical Considerations for the Review of*

Research Applications, assists reviewers in understanding the unique context of funding applications that propose research with Indigenous communities.³² Failing to follow these guidelines may produce inaccurate or difficult-to-interpret results and is also unethical. We hope the recommendations we have outlined here provide an introduction to alternative practices as public health researchers endeavor to represent data about Native communities with the dignity and nuance that all people deserve. *AJPH*

ABOUT THE AUTHORS

Danielle R. Gartner (Sault Ste. Marie Tribe of Chippewa Indians) is with the Department of Epidemiology & Biostatistics, College of Human Medicine, Michigan State University, East Lansing. Rachel E. Wilbur (Tolowa descent and Chetco descent) is with the Department of Anthropology, College of Arts & Sciences, and the Carolina Population Center, University of North Carolina, Chapel Hill. Meredith L. McCoy (Turtle Mountain Band of Chippewa Indians descent) is with the Department of American Studies and the Department of History, Carleton College, Northfield, MN.

CORRESPONDENCE

Correspondence should be sent to Danielle R. Gartner, PhD, Department of Epidemiology and Biostatistics, 909 Wilson Rd, B649 West Fee Hall, Michigan State University, East Lansing, MI 48824 (e-mail: gartnerd@msu.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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Trends in Cannabis Involvement and Risk of Alcohol Involvement in Motor Vehicle Crash Fatalities in the United States, 2000–2018

Marlene C. Lira, MPH, Timothy C. Heeren, PhD, Magdalena Buczek, MS, Jason G. Blanchette, JD, MPH, Rosanna Smart, PhD, Rosalie Liccardo Pacula, PhD, and Timothy S. Naimi, MD, MPH

Objectives. To assess cannabis and alcohol involvement among motor vehicle crash (MVC) fatalities in the United States.

Methods. In this repeated cross-sectional analysis, we used data from the Fatality Analysis Reporting System from 2000 to 2018. Fatalities were cannabis-involved if an involved driver tested positive for a cannabinoid and alcohol-involved based on the highest blood alcohol concentration (BAC) of an involved driver. Multinomial mixed-effects logistic regression models assessed cannabis as a risk factor for alcohol by BAC level.

Results. While trends in fatalities involving alcohol have remained stable, the percentage of fatalities involving cannabis and cannabis and alcohol increased from 9.0% in 2000 to 21.5% in 2018, and 4.8% in 2000 to 10.3% in 2018, respectively. In adjusted analyses, fatalities involving cannabis had 1.56 (95% confidence interval [CI] = 1.48, 1.65), 1.62 (95% CI = 1.52, 1.72), and 1.46 (95% CI = 1.42, 1.50) times the odds of involving BACs of 0.01% to 0.049%, 0.05% to 0.079%, and 0.08% or higher, respectively.

Conclusions. The percentage of fatalities involving cannabis and coinvolving cannabis and alcohol doubled from 2000 to 2018, and cannabis was associated with alcohol coinvolvement. Further research is warranted to understand cannabis- and alcohol-involved MVC fatalities. (*Am J Public Health.* 2021;111(11):1976–1985. <https://doi.org/10.2105/AJPH.2021.306466>)

Motor vehicle crashes (MVCs) are a leading cause of injury death in the United States, and more than 35% of MVC fatalities result from crashes in which at least 1 driver has a positive blood alcohol concentration (BAC).^{1,2} Cannabis use is a risk factor for driving impairment and the second-most-common substance involved in fatal MVCs after alcohol.^{3,4} The impairing effects of Δ 9-tetrahydrocannabinol (THC) in cannabis on driving ability include lane weaving, delayed reaction time, decreased coordination, and

distorted perception.^{4–9} Laboratory tests show that low-dose cannabis (< 10% THC) used in combination with alcohol may increase impairment more than either substance alone, particularly for skills relevant to driving.^{10–13}

Throughout the past 2 decades, cannabis policy in the United States has dramatically shifted as states have legalized medical cannabis, decriminalized possession of cannabis, and legalized recreational cannabis; correspondingly, rates of cannabis use in the United States have increased among

adults.¹⁴ Between 2002 and 2017, past-month cannabis use assessed through the National Survey on Drug Use and Health (NSDUH) increased among adults aged 26 years and older from 4.0% to 7.9%, and among adults aged 18 to 25 years from 17.3% to 22.1%.¹⁵ Among nighttime drivers, rates of cannabis prevalence in oral fluid also increased, from 8.6% in the 2007 National Roadside Survey to 12.6% in the 2013–2014 National Roadside Survey.^{16,17} More recently, a study utilizing NSDUH data from 2016 to

2018 found that 29.5% of cannabis users reported driving under the influence of cannabis.¹⁸ While the proportion of US crash deaths that involve alcohol has remained constant over time despite significant safety improvements in cars and car driving services,^{2,19} positive toxicology results for substances other than alcohol in fatally injured drivers has increased to approximately 30%.²⁰

Little consideration has been given to trends in co-use of alcohol and cannabis, in particular cannabis involvement at various alcohol levels in MVC deaths. Alcohol coinvolvement may be 1 explanation for the inconsistent findings of cannabis policy changes on cannabis-involved MVC fatalities, in addition to the difficulties in disentangling whether cannabis and alcohol are used as substitutes or complements, and whether changes in actual cannabis involvement stem from changes in testing for cannabis, given suboptimal testing.²¹⁻²⁷ Thus, there is still a need to describe cannabis involvement in MVC fatalities at various levels of alcohol involvement and to assess whether cannabis is a contributing risk factor for alcohol-involved crashes.

To fill this gap in the literature, the objectives of this study were 3-fold. First, we sought to describe drug testing rates in MVC fatalities overall and by BAC level. Second, we sought to describe patterns of alcohol and cannabis coinvolvement in MVC fatalities overall and by BAC level, and evaluated differences in decedent characteristics across crashes with varying substance involvement and coinvolvement. Finally, we assessed the cannabis involvement as a risk factor for alcohol coinvolvement by BAC level in MVC fatalities. Given the conflicting evidence regarding cannabis and alcohol as potential

substitutes or complements, we did not have prespecified hypotheses regarding the direction or magnitude of potential associations between cannabis and alcohol.

METHODS

The methods of this study were similar to those conducted by our study team in past analyses.^{2,19,28} This was a repeated cross-sectional analysis of MVC fatalities from the Fatality Analysis Reporting System (FARS), a census of MVC fatalities in US states and Washington, DC, from 2000 to 2018.²⁹

Participants

In this study, we utilized distinct participant groups for each stage of the analysis. For the first objective, describing drug testing rates in MVC fatalities overall and by BAC level, participants included FARS decedents from 2000 to 2018 who died in MVCs in which at least 1 driver was identified. For the second objective, describing patterns of alcohol and cannabis coinvolvement, participants were restricted to those with crash-level drug testing, defined as at least 1 driver having valid drug test results. For the third objective of the study, assessing cannabis involvement as a risk factor for alcohol coinvolvement, the group was further restricted to those with individual- and crash-level characteristics.

Variables

The main predictor, crash-level cannabis involvement, was defined as at least 1 driver with valid drug test results indicating the presence of a cannabinoid (FARS drug test results from 600 to 695). Before 2018, FARS reported up to

3 substances in drug test results based on a hierarchy, such that narcotics would be given top priority, followed by depressants, stimulants, hallucinogens, and then cannabinoids.³⁰ In 2018, FARS began reporting all positive substances. To use more recent data and maintain a uniform definition, the hierarchy used through 2017 was applied to data from 2018. The outcome, alcohol involvement by BAC level, was similarly calculated at the crash level such that the highest BAC from all drivers was assigned to all decedents from the crash. BAC levels were 0.00%, 0.01% to 0.049%, 0.05% to 0.079%, and 0.08% or higher. Because it is unlikely that alcohol test results are missing completely at random or missing completely not at random, missing results are an important threat to validity. Therefore, we utilized validated, imputed data sets from FARS to estimate missing BAC levels.³¹

Potential confounders were year (continuous); decedent-level characteristics: sex (male [reference] vs female), age category (< 21 years [reference], 21-34 years, 35-54 years, ≥ 55 years), race/ethnicity (White [reference], non-Hispanic Black, Hispanic, other, unknown); crash-level characteristics: opioid involvement (no [reference] vs yes), other substance involvement (i.e., ≥ 1 driver positive for substance other than alcohol, cannabis, or opioids; no [reference] vs yes), urbanicity (rural [reference] vs urban); and continuous state-level characteristics: percentage male, percentage non-Hispanic White, percentage Hispanic, percentage aged 21 years or older, percentage with college degree or higher, percentage Catholic population, median household income, law enforcement officers per 1000 residents, annual vehicle miles traveled, and state-year drug testing rate.³²⁻³⁵ Decedents with missing

individual- or crash-level characteristics were excluded from analyses provided that missingness was not substantial (< 1%). Approximately 9% of decedents had missing information on race/ethnicity, so an “unknown” category was created.

Statistical Methods

We calculated crash-level drug testing rates and cannabis involvement rates overall and by BAC level over the study period. Because testing rates changed over time, we also calculated rates of cannabis involvement after restricting the data to state-years with testing rates of at least 50%, restricting to states with testing rates of at least 50% for all years, and restricting to states with testing rates of at least 33.3% for all years. We calculated prevalence rates of cannabis involvement by demographic and crash-level characteristics in all MVC fatalities, as well as among various BAC strata.

We calculated the prevalence of cannabis by BAC level and demographic and crash characteristics. Because alcohol values were imputed for some decedents, it was possible that the level of alcohol involvement could change across imputations. The data set was therefore transformed from a wide format to a long format, such that there were 10 observations for each decedent. Cannabis involvement was tabulated overall and for each BAC level by the variable rows in Table 1. Values were then divided by 10 and rounded to the nearest whole number to return the data set to its original size.

We used mixed-effects multinomial logistic regression models with the multiply imputed alcohol data using Rubin's rules³⁶ to assess crude and adjusted relationships between crash-level

cannabis involvement and alcohol levels (BACs of 0.01%–0.049%, 0.05%–0.079%, and $\geq 0.08\%$ vs 0.00%). Given that there may be similarities within states, we accounted for potential clustering by state random effects. Covariates in adjusted analyses included those listed in the “Variables” section.

Sensitivity Analyses

Given suboptimal drug testing rates and the variation in approaches utilized by similar analyses in the wider literature, we performed a number of sensitivity analyses to assess the robustness of findings. These included further adjusting for the alcohol policy environment,³⁷ restricting the sample to only include decedents who were drivers, restricting the sample to only include decedents who were drivers and using individual-level toxicology results rather than crash-level toxicology results, conducting a crash-level analysis, utilizing nonimputed data, and restricting the sample to include decedents from state-years with testing rates of at least 50% and 66.67%. We conducted analyses by using Stata version 15.1 (Stata-Corp LP, College Station, TX).

Effect Modification

Given that age and sex are strong risk factors for cannabis and alcohol use, we conducted posthoc analyses stratified on the basis of decedent sex (male vs female) and age (< 21 years, 21–34 years, 35–54 years, and ≥ 55 years).

RESULTS

Between 2000 and 2018, there were 721 825 MVC fatalities in the United States with at least 1 identified driver. Of these, 327 073 (45.3%) had crash-level

drug testing results, constituting the sample for Figures 1 and 2. Of these, 322 773 (98.7%) had complete demographic and crash information, constituting the final sample for analyses (Tables 1 and 2). The 322 773 decedents comprised 254 002 drivers, 52 053 passengers, and 16 718 other victims (i.e., pedestrians, cyclists). Of the 254 002 deceased drivers, 243 926 or 96.0% were actually tested for alcohol. When alcohol results were applied to nondriver decedents at the crash level, 270 311 of the 322 773, or 83.8%, had at least 1 driver with valid alcohol test results.

Drug testing overall increased from 32.9% in 2000 to 47.9% in 2018 (data not shown). Overall and when stratified by crash-level BAC, testing increased for all BAC levels over time (Figure A, available as a supplement to the online version of this article at <http://www.ajph.org>); however, MVC fatalities not involving alcohol had the lowest rates of drug testing.

The percentage of fatalities involving cannabis increased from 9.0% in 2000 to 21.5% in 2018 (Figure 1). Given that testing and cannabis both increased during the study period, we also assessed cannabis involvement restricting the data to state-years with higher testing rates, and results were consistent (Figure B, available as a supplement to the online version of this article at <http://www.ajph.org>). After restricting state-years to those with testing rates of at least 50%, restricting the sample to states with testing rates of at least 50% for all years (n = 7: Hawaii, Illinois, New Hampshire, New Jersey, Rhode Island, Washington, and West Virginia), and restricting the sample to states with testing rates of at least 33.3% for all years (n = 16: the aforementioned states as well as California, Connecticut,

TABLE 1— Number and Percentage of Motor Vehicle Crash Fatalities That Were Cannabis-Involved, by Blood Alcohol Concentration (BAC) Level and Demographic and Crash Characteristics: Fatality Analysis Reporting System, United States, 2000–2018

Characteristic	Overall (n = 322 773), No. (%)	BAC 0.00% (n = 192 367), No. (%)	BAC 0.01%–0.049% (n = 12 182), No. (%)	BAC 0.05%–0.079% (n = 9146), No. (%)	BAC ≥ 0.08% (n = 109 078), No. (%)
Fatalities					
All fatalities	45 512 (14.1)	22 108 (11.5)	2 302 (18.9)	1 846 (20.2)	19 256 (17.7)
Drivers	34 048 (13.4)	16 047 (10.8)	1 596 (17.3)	1 224 (18.3)	15 181 (16.9)
Passengers	9 461 (18.2)	4 642 (15.1)	595 (24.5)	540 (26.0)	3 684 (21.9)
Others ^a	2 003 (12.0)	1 420 (10.4)	110 (21.5)	82 (22.6)	390 (18.4)
Sex					
Male	34 765 (14.9)	16 318 (12.5)	1 763 (19.3)	1 416 (20.2)	15 268 (17.7)
Female	10 747 (12.0)	5 790 (9.4)	539 (17.6)	430 (20.1)	3 988 (17.5)
Age, y					
< 21	9 469 (19.1)	5 437 (17.3)	528 (25.7)	427 (26.7)	3 077 (24.4)
21–34	18 939 (19.3)	7 858 (17.2)	921 (24.0)	813 (23.8)	9 348 (20.7)
35–54	12 011 (12.3)	5 593 (10.5)	573 (15.1)	441 (16.7)	5 403 (14.4)
≥ 55	5 093 (6.6)	3 221 (5.4)	280 (11.2)	165 (11.1)	1 427 (10.5)
Race/ethnicity					
Non-Hispanic White	29 151 (13.9)	14 410 (11.1)	1 410 (18.2)	1 159 (20.3)	12 172 (18.2)
Non-Hispanic Black	6 327 (18.7)	3 052 (16.3)	371 (23.4)	289 (24.5)	2 615 (21.2)
Hispanic	4 856 (12.8)	2 204 (11.4)	272 (18.8)	199 (16.9)	2 182 (13.6)
Other	1 605 (13.1)	701 (9.8)	68 (17.2)	63 (19.9)	774 (17.7)
Unknown	3 573 (12.5)	1 742 (10.0)	181 (18.1)	137 (17.8)	1 514 (15.9)
Involvement of other drugs					
No other drugs	30 672 (12.6)	14 235 (9.7)	1 391 (16.7)	1 202 (19.1)	13 843 (16.9)
Other drugs	14 840 (18.6)	7 873 (17.1)	910 (23.8)	644 (22.5)	5 412 (20.1)
Urbanicity					
Rural	24 139 (13.4)	12 093 (10.9)	1 221 (18.0)	959 (19.7)	9 866 (17.1)
Urban	21 373 (15.0)	10 016 (12.3)	1 080 (20.0)	888 (20.8)	9 390 (18.3)
No. of vehicles					
Single vehicle crash	23 959 (14.9)	9 695 (11.6)	1 137 (19.8)	1 044 (21.7)	12 084 (18.0)
Multiple vehicle crash	21 553 (13.3)	12 414 (11.4)	1 165 (18.1)	803 (18.5)	7 172 (17.0)
Day					
Weekday	24 910 (13.3)	14 426 (11.1)	1 162 (18.4)	851 (20.2)	8 471 (17.9)
Weekend	20 602 (15.2)	7 683 (12.2)	1 139 (19.4)	995 (20.2)	10 785 (17.4)
Time of day					
06:00–08:59	3 901 (12.1)	2 748 (10.6)	162 (16.9)	109 (20.6)	882 (18.2)
09:00–11:59	3 207 (10.3)	2 652 (9.8)	128 (15.5)	71 (17.8)	356 (12.8)
12:00–14:59	4 782 (11.3)	3 705 (10.5)	219 (15.4)	113 (16.4)	745 (14.5)
15:00–17:59	6 607 (12.6)	4 332 (11.3)	355 (17.1)	235 (17.5)	1 685 (15.6)
18:00–20:59	7 168 (14.8)	3 315 (12.6)	413 (19.2)	325 (19.2)	3 116 (17.0)
21:00–23:59	7 316 (16.3)	2 409 (13.5)	419 (21.3)	342 (20.3)	4 146 (17.8)
00:00–02:59	7 820 (18.3)	1 492 (15.1)	376 (22.6)	428 (24.5)	5 523 (18.8)
03:00–05:59	4 711 (16.6)	1 454 (12.4)	229 (20.7)	224 (21.1)	2 804 (19.3)

^aPedestrians, cyclists, etc.

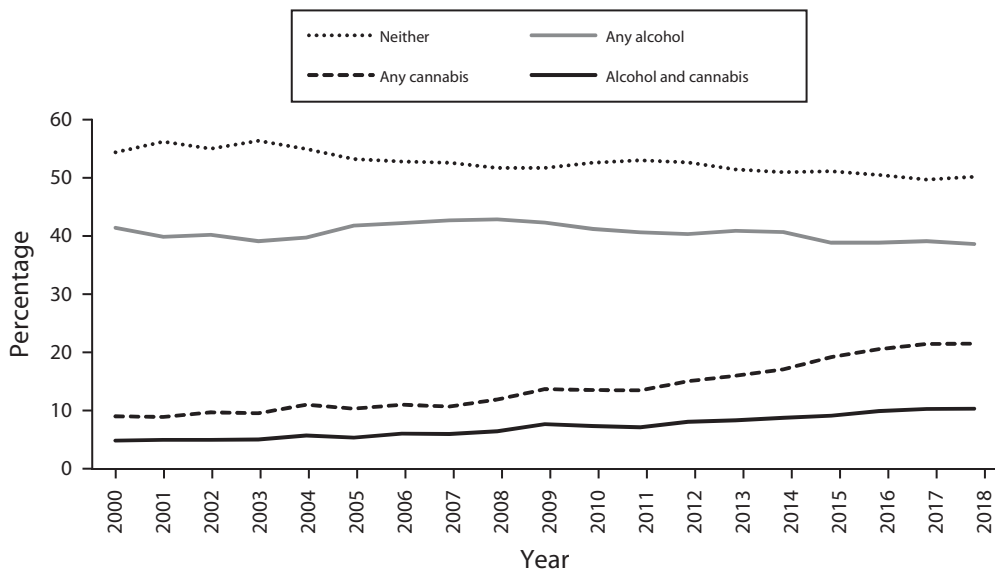


FIGURE 1— Percentage of Motor Vehicle Crash Fatalities With Crash-Level Drug Testing With Neither Alcohol nor Cannabis Involvement, Any Alcohol Involvement, Any Cannabis Involvement, or Both Alcohol and Cannabis Involvement: Fatality Analysis Reporting System, United States, 2000–2018

Note. Alcohol involvement defined as a crash-level blood alcohol concentration greater than 0.00%.

Georgia, Kentucky, Nevada, New Mexico, North Dakota, Ohio, and Wyoming), the rates of cannabis involvement were similar, indicating that changes in testing have not substantially affected

observed levels of cannabis involvement in fatal crashes.

We examined the percentage of decedents from crashes with any alcohol, any cannabis, cannabis and

alcohol, and neither alcohol nor cannabis over the study period (Figure 1). Alcohol was consistently involved in approximately 40% of MVC fatalities from 2000 to 2018. However, the

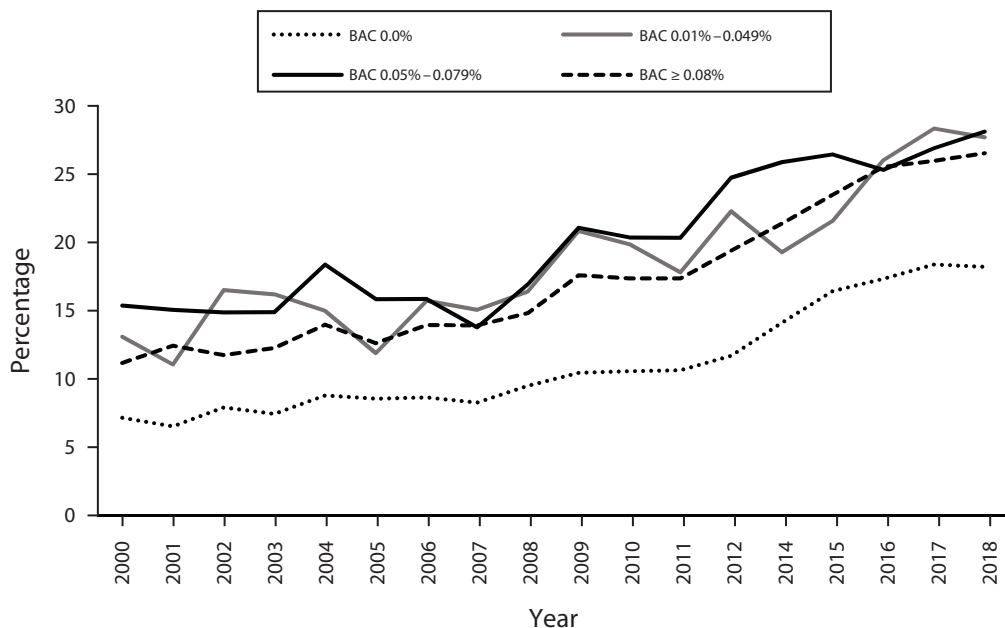


FIGURE 2— Percentage of Motor Vehicle Crash Fatalities Involving Cannabis by Blood Alcohol Concentration (BAC) Level: Fatality Analysis Reporting System, United States, 2000–2018

TABLE 2— Mixed Effects Multinomial Logistic Regression Models of Blood Alcohol Concentration (BAC) Level Based on Prevalence of Cannabis Involvement: Fatality Analysis Reporting System, United States, 2000–2018

Predictor: Cannabis Involvement	No.	Outcome, OR (95% CI) ^a		
		BAC 0.01%–0.049%	BAC 0.05%–0.079%	BAC ≥ 0.08%
Main analyses				
Overall unadjusted	322 773	1.77 (1.68, 1.86)	1.95 (1.84, 2.06)	1.67 (1.63, 1.71)
Overall adjusted ^b	322 773	1.56 (1.48, 1.65)	1.62 (1.52, 1.72)	1.46 (1.42, 1.50)
Stratified analyses				
Adjusted, stratified by sex ^c				
Men	232 899	1.49 (1.41, 1.59)	1.53 (1.43, 1.63)	1.38 (1.34, 1.42)
Women	89 874	1.84 (1.66, 2.04)	1.99 (1.76, 2.25)	1.78 (1.69, 1.86)
Adjusted, stratified by age, ^d y				
< 21	49 530	1.69 (1.52, 1.89)	1.83 (1.61, 2.08)	1.66 (1.57, 1.75)
21–34	98 262	1.49 (1.37, 1.62)	1.49 (1.36, 1.63)	1.29 (1.24, 1.34)
35–54	97 435	1.44 (1.30, 1.58)	1.62 (1.44, 1.82)	1.46 (1.40, 1.53)
≥ 55	77 546	2.06 (1.80, 2.36)	1.96 (1.63, 2.35)	1.96 (1.82, 2.10)
Sensitivity analyses				
Adjusted, ^b with addition of Alcohol Policy Scale ³⁸ score	322 773	1.56 (1.48, 1.65)	1.62 (1.52, 1.71)	1.46 (1.42, 1.50)
Adjusted, ^b restricted to drivers, crash-level test results	254 002	1.46 (1.36, 1.56)	1.48 (1.38, 1.59)	1.42 (1.39, 1.46)
Adjusted, ^e restricted to drivers, individual-level test results	243 528	1.45 (1.36, 1.55)	1.56 (1.45, 1.68)	1.52 (1.48, 1.56)
Adjusted, ^f crash-level	287 230	1.74 (1.65, 1.84)	1.93 (1.81, 2.05)	1.68 (1.64, 1.73)
Adjusted, ^b nonimputed data	270 311	1.45 (1.37, 1.54)	1.48 (1.38, 1.58)	1.41 (1.38, 1.45)
Adjusted, ^b restricted to state-years with testing rates ≥ 50%	211 336	1.58 (1.48, 1.69)	1.63 (1.52, 1.76)	1.48 (1.44, 1.53)
Adjusted, ^b restricted to state-years with testing rates ≥ 66.67%	65 887	1.68 (1.50, 1.88)	1.45 (1.26, 1.66)	1.52 (1.45, 1.60)

Note. CI = confidence interval; OR = odds ratio.

^aRef = BAC 0.00%.

^bAdjusted for year; decedent-level characteristics: sex (male [Ref] vs female), age category (< 21 y [Ref], 21–34 y, 35–54 y, ≥ 55 y), race/ethnicity (White [Ref], non-Hispanic Black, Hispanic, other, unknown); crash-level characteristics: opioid involvement (no [Ref] vs yes), other substance involvement (no [Ref] vs yes), urbanicity (rural [Ref] vs urban); and continuous state-level characteristics: percentage male, percentage non-Hispanic White, percentage Hispanic, percentage aged ≥ 21 y, percentage with college degree or higher, percentage Catholic population, median household income, law enforcement officers per 1000 residents, and state-year drug testing rate.

^cAdjusted for covariates in footnote a, except sex.

^dAdjusted for covariates in footnote a, except age.

^eAdjusted for covariates in footnote a, but using individual-level alcohol, cannabis, opioid, and other drug involvement.

^fCrash-level analyses adjusted for crash- and state-level characteristics but not decedent-level characteristics.

percentage of crashes involving any cannabis more than doubled from 9.0% in 2000 to 21.5% in 2018, and the percentage of fatalities involving both cannabis and alcohol more than doubled from 4.8% to 10.3%. When we examined cannabis involvement by BAC level (Figure 2), cannabis involvement increased over time for all groups. Cannabis involvement was

more prevalent among fatalities that involved alcohol at all 3 BAC-level categories compared with fatalities that were not alcohol-involved.

We examined cannabis involvement by decedent and crash characteristics overall (Table 1). Crash-level cannabis involvement was more prevalent among decedents who were younger than 35 years and who were non-

Hispanic Black, and in crashes that involved other drugs, occurred at night, and occurred on weekends. Decedents who were passengers were also more likely to have died in accidents involving cannabis, suggesting that these crashes are more likely to involve deaths of individuals other than the driver. Cannabis involvement by decedent and crash characteristics were also examined and

stratified by level of alcohol involvement. Cannabis involvement was more prevalent among decedents with alcohol involvement among all BAC strata than decedents without alcohol involvement. Specifically, cannabis was involved in 11.5% of alcohol-uninvolved fatalities, versus 18.9% of fatalities involving BACs of 0.01% to 0.049%, 20.2% of fatalities involving BACs from 0.05% to 0.079%, and 17.7% of fatalities with BACs of 0.08% or higher. Among decedents who were younger than 21 years, 17.3% of deaths from alcohol-uninvolved crashes involved cannabis, whereas approximately 25% of deaths from alcohol-involved crashes involved cannabis. Similarly, among decedents who died in evening crashes, 13.2% of alcohol-uninvolved deaths involved cannabis, but approximately 20% of alcohol-involved deaths also involved cannabis.

Regression Analyses

In unadjusted, multinomial, multiple imputation analyses, fatalities involving cannabis had 1.77 (95% confidence interval [CI] = 1.68, 1.86) times the odds of involving a BAC from 0.01% to 0.049%, 1.95 (95% CI = 1.84, 2.06) times the odds of involving a BAC from 0.05% to 0.079%, and 1.67 (95% CI = 1.63, 1.71) times the odds of involving a BAC of 0.08% or higher compared with fatalities not involving cannabis (Table 2). In models adjusted for year, decedent-level characteristics, crash-level characteristics, and state-level characteristics, fatalities involving cannabis had 1.56 (95% CI = 1.48, 1.65) times the odds of involving a BAC from 0.01% to 0.049%, 1.62 (95% CI = 1.52, 1.72) times the odds of involving a BAC from 0.05% to 0.079%, and 1.46 (95% CI = 1.42, 1.50) times the odds of

involving a BAC of 0.08% or higher compared with fatalities not involving cannabis. The full adjusted model can be found in Table B (available as a supplement to the online version of this article at <http://www.ajph.org>).

Additional Analyses

Results were similar across the aforementioned sensitivity analyses further adjusted for alcohol policies, restricted to drivers, at the crash-level, using non-imputed data, and restricted to fatalities from state-years with various thresholds for testing. Given that decedent sex and age were associated with alcohol involvement, we conducted posthoc analyses stratified by age and sex to assess whether the relationship between cannabis and alcohol involvement differed on the basis of these factors (Table 2). Although females, individuals younger than 21 years, and individuals aged 55 years or older had lower odds of alcohol involvement in the main analysis (Table B), the relationship between cannabis and alcohol involvement appeared stronger for these groups.

DISCUSSION

In this study, we assessed trends in cannabis and alcohol involvement in MVC decedents in the United States from 2000 to 2018 and examined how cannabis involvement relates to alcohol involvement. To our knowledge, this is the first study to examine recent trends over time of cannabis involvement in relation to alcohol involvement in the United States, and the first to examine these relationships among all crash decedents rather than only drivers. We found that cannabis involvement and cannabis and alcohol

involvement in fatal MVCs are increasing nationally. While rates of alcohol involvement have remained steady over time, the rates of cannabis involvement have increased. This does not offer support for the idea of cannabis and alcohol being substitutes, at least in terms of MVC fatalities. In adjusted regression analyses, cannabis was associated with alcohol involvement, even at BAC levels below 0.08%, indicating that cannabis use is a risk factor for alcohol-involved MVC fatalities even at levels of alcohol below the legally permissible level for driving.

The proportion of MVC fatalities that were cannabis-involved more than doubled during the study period. This could be attributable to shifting cannabis policies enabling expansion of medical and recreational cannabis markets, changing societal attitudes toward cannabis, and other factors such as increased cannabis potency. Although low drug testing rates are a known limitation of FARS, they increased from 32.9% to 47.9% over the study period. However, trends in increased cannabis involvement over time were consistent when we looked at subsets of states and state-years with higher testing rates, as has been done in past studies.³⁹ Drug testing procedures and rates are heterogeneous across states, and some states may conduct drug testing selectively, either opting for crashes in which alcohol use is already suspected, or the opposite, not conducting drug testing when alcohol is already known to be involved as a cost-saving measure, potentially resulting in biases in either direction.³⁹⁻⁴¹

Cannabis involvement was more prevalent among fatalities that involved alcohol than fatalities that did not involve alcohol (Table 1). These findings were remarkably consistent by

demographic and crash-level characteristics, with the highest prevalence of cannabis and alcohol coinvolvement in crashes involving younger decedents (aged < 35 years) and passenger deaths. Although most alcohol-involved fatalities occur at BACs of 0.08% or higher, the legal limit in all states but Utah (which in 2019 implemented a limit of 0.05%), cannabis coinvolvement was similar across BAC levels in crashes involving any alcohol. In regression analyses, cannabis involvement was associated with increased odds of also involving alcohol, regardless of BAC level. This finding was consistent across multiple sensitivity analyses and suggests that cannabis is a risk factor for alcohol-involved MVCs.

This study builds on past epidemiological studies of alcohol- and cannabis-involved MVC fatalities. While laboratory studies have shown driving impairment from cannabis and synergistic effects with alcohol, only a handful of recent epidemiological studies have examined the combined effects of alcohol and cannabis on the risk of crashes. Two case-control studies found a dose-effect of alcohol and cannabis on unsafe driving actions,^{42,43} and another study found that culpable drivers in fatal MVCs were more likely to test positive for alcohol, cannabis, or both.⁴⁴ A meta-analysis conducted by the same authors found a pooled effect of marijuana on crash risk to be 2.66 (95% CI = 2.07, 3.41).⁴⁵ Contrasting results were seen in 2 studies that compared fatally injured drivers from FARS with matched drivers from the National Roadside Survey: one reported marijuana being associated with 83% increased odds of being a fatally injured driver rather than a control,⁴⁶ and the other reported statistically nonsignificant findings regardless of alcohol involvement.⁴⁷ These diverging

results were later attributed to methodological differences, such as inclusion factors for states based on testing rates, and the authors recommended future research should “account for as many factors as possible when assessing crash risk,” which we have tried to do in this analysis.^{39(p324)} More precise estimates of the degree of cannabis involvement, and the nature of alcohol-cannabis involved crash fatalities, would require testing levels of cannabinoids in drivers.

Given that the percentage of MVC fatalities involving alcohol has remained relatively stable at approximately 40% over the past 2 decades, it could be that increases in coinvolvement of cannabis are undercutting attempts at reducing alcohol-involved crash fatalities. There is a known relationship between alcohol policies and alcohol involvement in MVCs among adults and children, and within crashes involving alcohol at levels below the legal limit of 0.08%.^{2,19,28} However, future research is warranted to understand whether cannabis changes the protective associations between alcohol policies and alcohol-involved crashes.

A number of studies have analyzed the effect of various forms of cannabis legalization on cannabis-involved MVC fatalities; however, results have been conflicting, and these studies have not consistently examined interactions with alcohol or alcohol policies.²¹⁻²⁷ Adopting a lower permissible BAC threshold for those with cannabis in their system may be a policy strategy to reduce MVC harms from concurrent and simultaneous use of alcohol and cannabis. Indeed, even without consideration for cannabis, the National Highway Traffic Safety Administration and the National Academies of Sciences, Engineering, and Medicine have recommended

decreasing the legal alcohol limit to 0.05% to reduce alcohol-involved MVC fatalities.⁴⁸

Limitations

This study is subject to a number of limitations. Most importantly, these data cannot provide information on whether cannabis caused the crash. The mere presence of cannabis without a level is not alone indicative of impairment as, depending on frequency and amount of use, individuals can test positive for days or even weeks after use. Cannabinoid test results were collapsed such that individual cannabinoids were not analyzed. Previous research has suggested a blood THC limit of 5 nanograms per milliliter as a cutoff for impairment; however, FARS reports binary results.^{7,8} There has been variation in drug testing practices between states and over time, and some states may systematically not test for cannabis. This would, however, likely lead to an underestimation of the prevalence and bias toward the null in regression results. In addition, there is some evidence to suggest that simultaneous alcohol use increases THC levels and that there are lingering impairment effects once THC levels have declined, further complicating this.¹⁰

Although we conducted sensitivity analyses to understand the effect of limited testing on the prevalence of cannabis, it is possible that cannabis positivity was nevertheless subject to testing bias. In addition, because only 1 valid drug test from drivers in a crash was required for the definition of cannabis involvement, it is possible that cannabis involvement was undercounted if one driver tested negative and another was not tested. However, this would have likely resulted in a bias

toward the null, and the sensitivity analysis restricted to drivers and involving their individual-level test results mitigates this concern. Alcohol testing was similarly not uniform, and we used imputed data from FARS to account for suboptimal testing.³¹ Our regression analyses assessed the odds of alcohol involvement, to which the counterfactual is a fatality without alcohol involvement. This is distinct from incidence rate ratios but allows for adjusting for individual-level characteristics in models. However, our figures assessing rates over time support similar conclusions (Figure B).

Public Health Implications

Between 2000 and 2018, the percentage of MVC fatalities involving cannabis and involving cannabis and alcohol more than doubled. Fatalities involving cannabis had increased odds of also involving alcohol, regardless of BAC level, suggesting that cannabis use is a risk factor for alcohol-involved MVC fatalities in the United States. Future research is needed to understand how cannabis and alcohol relate to cannabis and alcohol policies, both separately and together. *AJPH*

ABOUT THE AUTHORS

Marlene C. Lira is with the Clinical Addiction Research and Education Unit, Department of Medicine, Boston Medical Center, Boston, MA. Timothy C. Heeren is with the Department of Biostatistics, Boston University School of Public Health, Boston. Magdalena Buczek is with the Boston University School of Medicine, Boston. Jason G. Blanchette is with the Department of Health Law, Policy, and Management, Boston University School of Public Health. Rosanna Smart is with RAND Corporation, Santa Monica, CA. Rosalie Liccardo Pacula is with the Department of Health Policy and Management, Sol Price School of Public Policy and Schaeffer Center for Health Policy and Economics, University of Southern California, Los Angeles. Timothy S. Naimi is with the Canadian Institute for Substance Use Research, University of Victoria, Victoria, BC.

CORRESPONDENCE

Correspondence should be sent to Marlene C. Lira, MPH, Clinical Addiction Research and Education Unit, Boston Medical Center, 801 Massachusetts Ave, 2nd Floor, Boston, MA 02118 (e-mail: marlene.lira@bmc.org). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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M. C. Lira, M. Buczek, and T. S. Naimi conceptualized the study. M. C. Lira conducted the analyses with statistical expertise provided by T. C. Heeren. All authors contributed to interpretation of the analysis. M. C. Lira drafted the article, and all authors provided critical feedback.

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CONFLICTS OF INTEREST

The authors have no conflicts of interest to report.

HUMAN PARTICIPANT PROTECTION

This study was determined to be not human participant research by the institutional review board at Boston University Medical Campus (protocol H-37378).

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Philadelphia's Excise Tax on Sugar-Sweetened and Artificially Sweetened Beverages and Supplemental Nutrition Assistance Program Benefit Redemption

Benjamin W. Chrisinger, PhD, MUEP

 See also Knox, p. 1907.

Objectives. To assess the effect of a 2017 excise tax on sugar and artificially sweetened beverages in Philadelphia, Pennsylvania, on the shopping patterns of low-income populations using Supplemental Nutrition Assistance Program (SNAP) data.

Methods. I used a synthetic controls approach to estimate the effect of the tax on Philadelphia and neighboring Pennsylvania counties (Bucks, Delaware, and Montgomery) as measured by total SNAP sales ("SNAP redemption") and SNAP redemption per SNAP participant. I assembled biannual data (2005–2019) from all US counties for SNAP redemption and relevant predictors. I performed placebo tests to estimate statistically significant effects and conducted robustness checks.

Results. Detectable increases in SNAP spending occurred in all 3 Philadelphia neighboring counties. Per-participant SNAP spending increased in 2 of the neighboring counties and decreased in Philadelphia. These effects were robust across multiple specifications and placebo tests.

Conclusions. The tax contributed to increased SNAP shopping in Philadelphia's neighboring counties across both outcome measures, and decreased spending in Philadelphia (at least by 1 measure). This raises questions about retailer behavior, the effectiveness of the tax's public health aim of reducing sugar-sweetened beverage consumption, and policy aims of investing in low-income communities. (*Am J Public Health.* 2021;111(11):1986–1996. <https://doi.org/10.2105/AJPH.2021.306464>)

Policies aimed at reducing population-level intake of sugar-sweetened beverages (SSBs) via excise taxes have now been adopted by several cities in the United States (e.g., Berkeley, Oakland, and San Francisco, California).¹ Recent evaluations of a \$.015-per-ounce excise tax in Philadelphia, Pennsylvania, a broader tax on both sugar-sweetened and artificially sweetened beverages, have

documented significantly higher prices and lower sales of taxed items following the policy's implementation on January 2017.² Though other research has not found any significant changes in unemployment claims across potentially affected industries, retailers and industry groups contend that recent store closures are directly related to the Philadelphia tax.³ Tax opponents often cite the regressive nature of the tax (i.e., on

average, lower-income individuals spend a higher proportion of their incomes on SSBs than higher-income individuals), and supporters highlight the tax's progressivity via its revenue-raising utility for antipoverty initiatives (e.g., universal prekindergarten schooling), a deliberate political feature of the Philadelphia tax identified by qualitative researchers.^{4,5} Still, to my knowledge, no research has investigated how

low-income shoppers responded to the beverage tax.

The socioeconomic context of food shopping is critical in Philadelphia. While the city's unemployment rate has declined from a 2012 peak of 10.9% to 6.2% in 2017 (vs national rates of 9.0% and 4.4%, respectively), food insecurity has risen over recent years, and the poverty rate has remained relatively flat at about 26%, the highest among the nation's largest 10 cities.^{6,7} In 2017, 22.1% of Philadelphia households participated in the federal government's largest and most widespread effort to reduce food insecurity, the Supplemental Nutrition Assistance Program (SNAP), nearly double the national rate of participation (11.2%).⁸ However, the average monthly amount of SNAP benefits redeemed at eligible retailers in Philadelphia also declined by \$5.4 million between 2016 and 2017, mirroring, at least partially, national declines in SNAP participation after a period of expansion following the 2008 economic crisis as part of the American Recovery and Reinvestment Act. During the same time, however, SNAP spending increased in the counties immediately neighboring Philadelphia, contrary to state and national trends. A key aim of this study was to identify approximately how much of these changes in SNAP spending, if any, is attributable to the tax.

Generally, SNAP is allocated monthly to participating households via electronic benefit transfers in amounts based on characteristics such as income, age, and the presence of children; nationwide, the average SNAP participant received about \$127 per month in 2018. The program is "cash-like," in that participants can redeem benefits on any food item (excluding prepared foods, alcohol, medicines, or

vitamins) at approved retailers. For many low-income households, SNAP benefits comprise a significant portion of monthly food budgets and are fully exhausted by the end of the month, though unused benefits are carried forward.^{9,10} While the program helps guard against experiences of deep poverty, food insecurity, and hospitalization,^{11,12} the nutritional content of SNAP purchases has also been the subject of scrutiny. Proposals from the political left and right call for restricting or excluding certain food items from SNAP,^{13,14} responding in part to studies that identify socioeconomic gradients in diet quality, with poorer households purchasing and consuming less healthy, often lower-cost foods.^{15,16} Cyclical patterns are commonly observed in analyses of benefit use, with end-of-month periods characterized by decreased food intake and lower nutritional quality of foods consumed.^{9,17} The program's relationship with neighborhood food environment is also notable: previous research has shown associations between SNAP retailer availability and participant enrollment,^{18,19} a multiplier effect of SNAP spending that stimulates broader economic activities,²⁰ and the monthly cycle in SNAP spending cited by retailers as a barrier to operating stores in lower-income communities.²¹ Thus, SNAP spending is key to our understanding of how policies might affect the health and everyday lives of socioeconomically disadvantaged groups.

METHODS

Monthly totals (2005–2019) for the dollar value of all SNAP benefits spent at eligible retailers in a given county (subsequently called "SNAP redemption") were obtained from the US

Department of Agriculture (USDA) for all counties in the United States from 2005 to 2019. Monthly counts (2005–2019) of stores where SNAP could be redeemed were also included. Redemption amounts from counties with 4 or fewer SNAP-eligible retailers are redacted by the USDA to protect retailer identities, though this typically only applies to very small counties.

The count of SNAP participants and amounts of SNAP benefits distributed were also obtained for every county from USDA; these data are released biannually, with month estimates provided for every January and July. State-level agencies that administer SNAP locally collect this information and provide it to USDA. Monthly county-level unemployment rate estimates were obtained from the US Bureau of Labor Statistics. County-level population estimates from the 2012–2016 5-year American Community Survey were also included as covariates. Table A (available as a supplement to the online version of this article at <http://www.ajph.org>) provides a summary of data sources.

Study Dates

To allow this analysis to incorporate government stimulus-related SNAP spending patterns, the starting month was selected as January 2005, several years before the onset of the American Recovery and Reinvestment Act. To temporally match data sets, only observations from January and July were included in this analysis, as this was the frequency of USDA reporting for county-level SNAP-participating individuals and SNAP benefits distributed. January 2019 observations were omitted from this analysis, as an anomalous spike appears in SNAP redemption

across all counties. July 2019 was selected as the end month, and the resulting monthly data set included 24 pretax observations and 5 posttax observations for each county.

Outcome Variables

The primary outcome variable was total county-level monthly SNAP redemption. As a secondary outcome, I calculated the value of SNAP redemption in a county per SNAP participant (hereafter called “SNAP per participant”) by dividing SNAP redemption by the count of SNAP participants in a county for a given month. I repeated the analyses subsequently described separately for each outcome.

Statistical Analyses

I employed a synthetic control approach to model the effect of the excise tax. I fit 4 separate synthetic control models (Philadelphia and its immediate neighbors, Bucks, Delaware, and Montgomery counties) for each outcome. Philadelphia is a consolidated city-county, so references to “Philadelphia,” “Philadelphia County,” or “city of Philadelphia” are interchangeable. I did not assess neighboring counties in New Jersey, given the additional physical and economic obstacles to travel (e.g., Delaware River with only toll-bridge crossings; see Figure A, available as a supplement to the online version of this article at <http://www.ajph.org>). This synthetic control method, pioneered by the work of Abadie, Gardeazabal, Diamond, and Hainmueller,^{22–24} has been used in previous social science research and is built upon interactive fixed-effect models, which generate hypothetical or “synthetic” controls for each treated unit, rather than a more

traditional matching approaches often used in evaluation research.^{24–26} The synthetic control approach leverages time-series outcomes across many potential donor units to identify optimal weights for estimating a counterfactual unit and can also integrate and weight the influence of covariates. Synthetic control studies typically employ a test statistic proposed by Abadie et al.—ratio of the root mean squared prediction errors (RMSPEs)—to compare models and calculate a treatment *P* value from a placebo test, which refits a synthetic control for each of the donor units.²⁴

Selection of donor counties. As described by Bouttelle et al., synthetic control methods rest on several assumptions: (1) the similarity of treated and potential control units, (2) no spillover effects of intervention to potential controls, and (3) no external shocks.^{27(p676)} To satisfy the first assumption, following the guidance of McClelland and Gault,²⁸ the pretreatment trends for SNAP redemption were also used to narrow the pool of potential donor counties. I calculated simple linear regressions for each potential donor county for 2 periods, 2008–2013 and 2014–2016, and extracted a slope to approximate the trend in SNAP redemption for the county. I chose these periods to reflect the expansion in SNAP under the American Recovery and Reinvestment Act, with peak SNAP enrollments occurring in 2013. Nationally, SNAP enrollment was declining in the second pretreatment period (2014–2016). At least 20 possible donor counties were selected based on their pretreatment predictor values and slopes.

This analysis directly addressed the second assumption, as it attempts to

measure the spillover of the Philadelphia policy into its most geographically accessible neighbors by fitting separate models for each neighboring county. These neighboring counties were excluded from each other’s potential donor pools (e.g., Montgomery County was excluded as a potential donor for synthetic Bucks and Delaware counties). Furthermore, any counties with similar local beverage taxes during the study period (San Francisco County and Alameda County in California)²⁹ and their immediate neighbors were also excluded from the data set. I used visual inspection of time-series outcome variable plots to identify potential external shocks (assumption 3).²⁶ I included only counties with complete data during the study period in the final data set.

Synthetic control estimation and specification searching. Recent methodological literature provides further guidance for systematic application of synthetic controls, especially as to the use of pretreatment lagged outcome observations in developing the synthetic unit.³⁰ The guidance provided by Ferman et al. is meant to offer a systematic way of specifying, choosing among, and testing different synthetic control specifications (Table B, available as a supplement to the online version of this article at <http://www.ajph.org>) for significant treatment effects, as a means of guarding against specification searching for statistical significance, and over- and underrejection of the null hypothesis.³⁰ In short, I used the following steps to identify synthetic control specifications and assess importance of the results: (1) visual inspection of pre-treatment gaps, which should be relatively small for well-fitting synthetic controls; (2) a *P* value less than .10 from the placebo test described by Abadie et al.,²⁴

(3) consistency of treatment effect observed across 95% confidence set calculations that combine specifications not rejected by steps 1 and 2; and (4) for the specification selected by a MSPE criterion proposed by Dube and Zipperer,³¹ consistency of the treatment effect observed in different covariate configurations, and “leave-one-out” and “in-time” placebo tests. These steps are described in greater detail in the Appendix (available as a supplement to the online version of this article). The assembled data set and code used for conducting the analyses are available on the Oxford University Research Archive (<https://doi.org/10.5287/odleian:0oqGkDBdy>).

RESULTS

The average monthly SNAP benefits redeemed in Philadelphia County in 2016 was \$61.78 million (SD = \$.8 million), and \$7.85 million across the 3 bordering counties in Pennsylvania, all higher than the monthly average for Pennsylvania counties (\$2.22 million) or the national average (\$1.8 million). In 2017, SNAP redemption fell by 6.5% in Philadelphia (to \$57.75 million), 1.9% in other Pennsylvania counties (to \$2.18 million), and 2.1% nationwide (to \$1.76 million). All of Philadelphia's neighboring Pennsylvania counties saw a rise in SNAP redemption: 3.1% in Bucks (to \$5.25 million), 6.0% in Montgomery (to \$9.77 million), and 8.9% in Delaware County (to \$10.06 million). [Figure 1](#) illustrates the trend for SNAP outcomes in Philadelphia and its immediate neighbors.

Over the same period, SNAP participation decreased nationally (−2.8%), with a smaller decrease in Pennsylvania counties (−0.7%). Philadelphia saw a decrease in SNAP participation of 1.4%,

while 2 of its neighbors saw steeper decreases (−3.9% in Montgomery and −2.2% in Bucks), and there was an increase of less than half a percent (.2%) in Delaware County. From 2016 to 2017, SNAP benefit distributions decreased by 4.4% nationwide, 1.4% in Pennsylvania counties, and 2.1% in Philadelphia. Among neighboring counties, decreases were also observed: −2.7% in Bucks, −0.5% in Delaware, and −5.3% in Montgomery.

SNAP Redemption Outcome

For SNAP redemption in Philadelphia, no specifications appeared to have a good fit, indicating that a synthetic control estimation may not be appropriate. Visual inspection of gap plots illustrated that the differences observed after treatment were not much larger than gaps seen between the treated and synthetic Philadelphia data in pretreatment periods. This may be reflective of the need to relax selection criteria to achieve a donor pool size larger than 20, indicating the difficulty in establishing a reasonable synthetic unit given the predictors used.

For SNAP redemption in Bucks County, all specifications indicated a positive long-run trend; of these, specification 5 (all odd pretreatment periods) had the lowest RMSPE ([Table 1](#)) and was also the selected specification per the criterion outlined by Dube and Zipperer.³¹ [Figure C](#) (available as a supplement to the online version of this article at <http://www.ajph.org>) illustrates relatively minor differences between specifications. Two covariates, total population and the number of authorized retailers, had marginal weights in the selected model (.022 and .001, respectively) ([Table 2](#)).

Placebo testing yielded a *P* value of .087 ([Figure D](#), available as a supplement to the online version of this article at <http://www.ajph.org>). Confidence sets for the effect functions of all specifications indicated a positive effect in the posttreatment period, with the earliest posttreatment periods crossing zero ([Figure E](#), available as a supplement to the online version of this article at <http://www.ajph.org>). Donor county weights are reported for all specifications in [Table C](#) (available as a supplement to the online version of this article at <http://www.ajph.org>). [Figure 2](#) summarizes the estimated gaps between treated and synthetic counties across the study period.

For Delaware County, all specifications also indicated a positive long-run effect, with specification 2 (first three quarters of pretreatment periods) selected based on the lowest RMSPE, though all specifications had similar RMSPE values; specification 5 was selected by the Dube and Zipperer criterion,³¹ though its outcomes and weightings were comparable to specification 2. Three covariates—the number of authorized stores, total SNAP benefits distributed, and number of SNAP participants—were marginally weighted in the model. Placebo testing produced a *P* value of .04, and confidence set plots also illustrated positive effects in all posttreatment periods.

For Montgomery County, all specifications also indicated a positive effect, and specification 3 (first half of pretreatment periods) was selected for the lowest RMSPE; specification 5 was selected by the Dube and Zipperer criterion,³¹ and produced similar though slightly larger gaps in posttreatment periods than specification 3. All covariates contributed at least marginal weights to the synthetic unit.

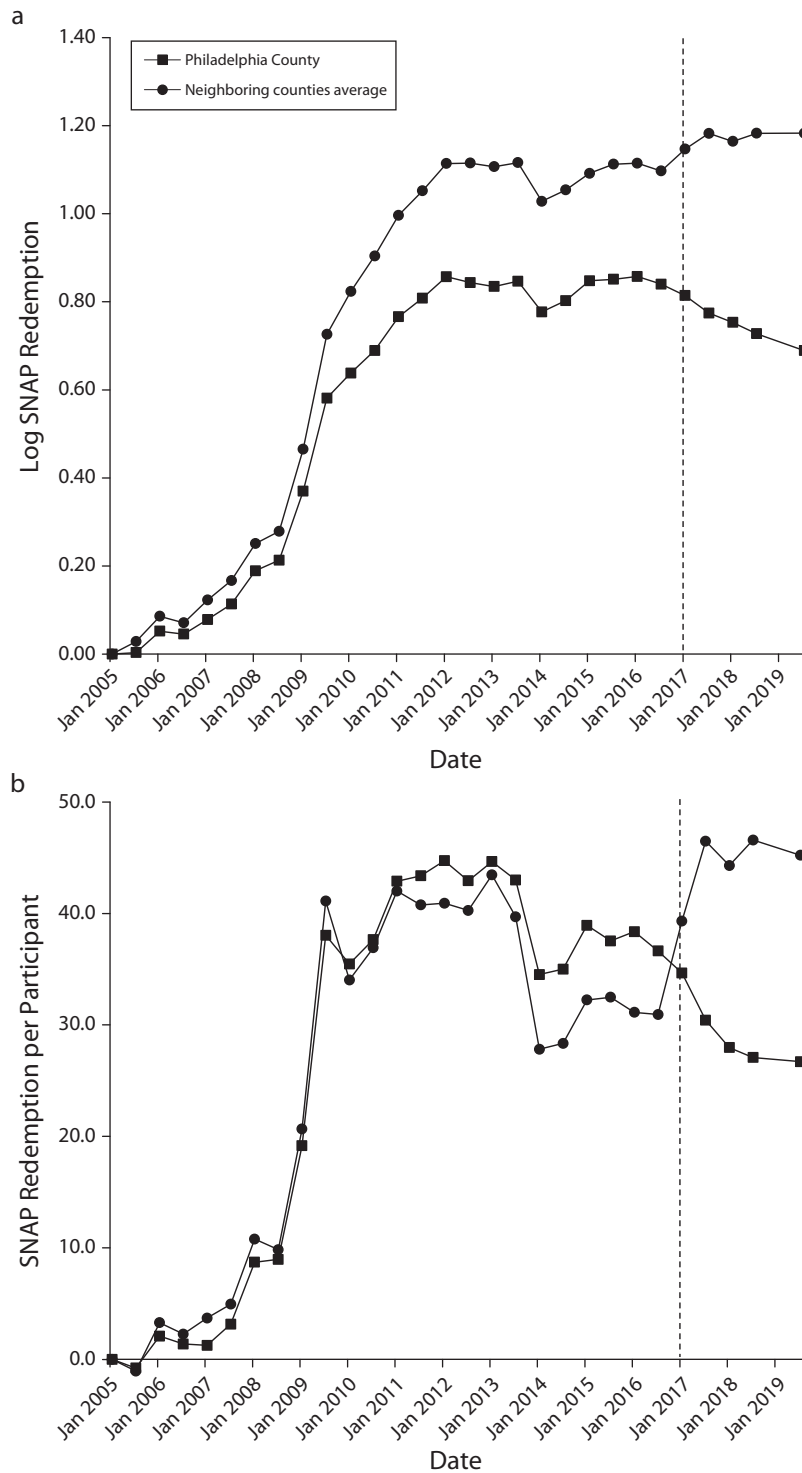


FIGURE 1— Monthly SNAP Outcomes in Philadelphia and Neighboring Pennsylvania Counties by (a) Log SNAP Redemption and (b) SNAP Redemption per Participant: 2005–2019, Indexed to January 2005

Note. SNAP = Supplemental Nutrition Assistance Program. Dashed vertical line indicates the implementation of the Philadelphia tax in January 2017. Neighboring counties include Bucks, Delaware, and Montgomery counties in Pennsylvania. An alternate version of this figure is provided as Figure B (available as a supplement to the online version of this article at <http://www.ajph.org>) and illustrates trends for each county separately.

Placebo testing produced a P value of .045, and confidence set plots showed a positive effect in all posttreatment periods.

Redemption Per Participant Outcome

For SNAP redemption amount per participant in Philadelphia, all specifications showed a negative long-run trend (Figure F, available as a supplement to the online version of this article at <http://www.ajph.org>), and 4 of 5 produced P values less than .05 via placebo testing (Figure G, available as a supplement to the online version of this article at <http://www.ajph.org>). Specification 5 had the smallest RMSPE (Table 1), with only the population covariate receiving a small weighting, reflected in the similarities of synthetic controls estimations among all 5 specifications (Table D, available as a supplement to the online version of this article at <http://www.ajph.org>). Specification 4 (odd pretreatment periods) was indicated by the selection criteria of Dube and Zipperer,³¹ though outcomes were similar to specification 5 (Figure F). Confidence sets for significant specifications illustrated a negative effect for most posttreatment periods (Figure H, available as a supplement to the online version of this article at <http://www.ajph.org>). Donor county weightings are reported in Table E (available as a supplement to the online version of this article at <http://www.ajph.org>).

All specifications for neighboring counties demonstrated a positive posttreatment trend. Specifications for Delaware County were deemed significant via placebo testing ($P = .029$ for all specifications in Delaware County). Specifications in Montgomery and

TABLE 1— Fit Criteria Used in Specification Searching for Synthetic Controls: Philadelphia, PA, and Neighboring Pennsylvania Counties, 2005–2019

Specification	Philadelphia County			Bucks County			Delaware County			Montgomery County		
	RMSPE	MSPE	P	RMSPE	MSPE	P	RMSPE	MSPE	P	RMSPE	MSPE	P
Log SNAP redemption												
All lags	0.849	0.002	.96	6.803	0.021	.09	6.456	0.032	.04	7.738	0.062	.045
First three quarters	1.275	0.006	.84	6.345	0.019	.09	6.458	0.034	.04	7.644	0.069	.045
First half	1.374	0.007	.84	6.643	0.020	.043	6.189	0.034	.04	6.584	0.052	.045
Odd lags	0.849	0.002	.96	6.814	0.021	.09	6.442	0.035	.04	7.502	0.058	.045
Even lags	0.849	0.002	.96	6.034	0.018	.09	6.222	0.048	.04	7.735	0.062	.045
SNAP redemption per participant												
All lags	6.520	33.51	.04	2.313	58.18	.13	9.485	288.24	.029	5.243	1048.11	.054
First three quarters	6.500	33.36	.038	2.331	59.12	.2	9.673	317.78	.029	5.244	1048.49	.054
First half	1.922	18.17	.46	2.952	112.15	.2	7.415	293.81	.029	5.242	1047.69	.054
Odd lags	6.510	33.38	.038	2.357	60.52	.24	9.530	292.06	.029	5.244	1048.24	.054
Even lags	6.462	33.02	.038	2.148	51.77	.28	9.222	295.56	.029	5.075	991.37	.054

Note. MSPE = mean squared prediction error; RMSPE = root mean squared prediction error; SNAP = Supplemental Nutrition Assistance Program. P values were calculated via placebo test method that reflects the rank of posttreatment gaps in the county under observation compared with placebo counties, divided by the number of placebo counties plus 1.

Source. Used approach described by Ferman et al.³⁰

Bucks counties did not yield significant P values through placebo testing ($P = .054$ for all specifications in Montgomery; specifications ranged from $P = .13$ to $.28$ in Bucks). In Delaware County, specification 3 had the lowest RMSPE, with SNAP benefit allocations, participant counts, and retailer counts included as covariates. In Montgomery County, specification 5 had the lowest RMSPE, with all covariates receiving at least a marginal weighting. Dube and Zipperer criteria³¹ indicated specification 5 for both counties. Confidence sets calculated for both Delaware and Montgomery counties illustrated a consistently positive trend across posttreatment periods.

DISCUSSION

Even amid broader macroeconomic changes, such as rates of SNAP participation, unemployment, and benefit allocations, implementation of the Philadelphia excise tax appears to have resulted in a geographical shift in SNAP spending. Though benefits cannot be directly traced from one county to another, synthetic control analyses for the 3 neighboring counties provides a high-level perspective on where these funds may have moved. Each neighbor saw detectable increases in SNAP redemption, with the largest increases observed in Montgomery County, which shares the longest border with Philadelphia of the 3 adjacent Pennsylvania counties. Similarly, per-participant SNAP redemption was shown to decrease in Philadelphia and increase in both Delaware and Montgomery counties. Here, it is reasonable to surmise that the observed changes in SNAP spending in Philadelphia's neighboring

TABLE 2— Treated and Synthetic Neighboring County Predictor Means and Weights for Log SNAP Redemption Outcome: Pennsylvania Counties Neighboring Philadelphia, 2005–2019

Predictor	Bucks County			Delaware County			Montgomery County		
	Treated Mean	Synthetic Mean	Weight	Treated Mean	Synthetic Mean	Weight	Treated Mean	Synthetic Mean	Weight
Population ^a	13.35	13.28	0.022	13.24	13.03	0.000	13.61	13.53	0.018
SNAP benefits ^a	14.95	14.86	0.000	15.65	15.79	0.004	15.2	15.45	0.001
SNAP participants ^a	10.22	10.17	0.000	10.89	11.04	0.009	10.46	10.66	0.017
SNAP retailers ^a	5.38	5.29	0.001	5.81	5.81	0.157	5.66	5.6	0.001
Unemployment rate	6	5.92	0.000	6.35	7.28	0.000	5.4	6.17	0.004
Redemption amount—lagged outcomes ^a									
Jan 2005	NA	NA	NA	15.11	15.13	0.066	14.77	14.74	0.066
Jul 2005	14.33	14.33	0.000	15.14	15.16	0.060	14.79	14.79	0.016
Jan 2006	NA	NA	NA	15.19	15.23	0.031	14.84	14.88	0.046
Jul 2006	14.37	14.37	0.216	15.19	15.18	0.059	14.83	14.85	0.189
Jan 2007	NA	NA	NA	15.23	15.2	0.097	14.9	14.88	0.006
Jul 2007	14.47	14.44	0.166	15.25	15.2	0.086	14.95	14.93	0.028
Jan 2008	NA	NA	NA	15.310	15.27	0.085	15.06	15.03	0.219
Jul 2008	14.59	14.57	0.062	15.31	15.31	0.065	15.11	15.09	0.230
Jan 2009	NA	NA	NA	15.47	15.49	0.049	15.31	15.33	0.002
Jul 2009	15.06	15.1	0.027	15.73	15.73	0.048	15.57	15.6	0.003
Jan 2010	NA	NA	NA	15.79	15.8	0.074	15.66	15.7	0.113
Jul 2010	15.3	15.31	0.197	15.85	15.89	0.110	15.73	15.78	0.040
Jan 2011	NA	NA	NA	NA	NA	NA	NA	NA	NA
Jul 2011	15.48	15.45	0.000	NA	NA	NA	NA	NA	NA
Jan 2012	NA	NA	NA	NA	NA	NA	NA	NA	NA
Jul 2012	15.53	15.51	0.027	NA	NA	NA	NA	NA	NA
Jul 2013	15.5	15.52	0.09	NA	NA	NA	NA	NA	NA
Jul 2014	15.45	15.44	0.03	NA	NA	NA	NA	NA	NA
Jul 2015	15.44	15.43	0.007	NA	NA	NA	NA	NA	NA
Jul 2016	15.43	15.44	0.154	NA	NA	NA	NA	NA	NA

Note. NA = specifications that do not include all observations; SNAP = Supplemental Nutrition Assistance Program.

^aLog-transformed.

counties is attributable to the implementation of the SSB tax. The mechanisms explaining these changes must be carefully considered.

Probably More Than Sugar-Sweetened Beverages

Earlier research has shown that decreases in Philadelphia SSB sales following the beverage tax were partially offset by increased sales in bordering counties, tempering expectations of broader public health benefits that would result from reduced SSB consumption.² This is consistent with other research from Oakland, California, which finds similar cross-border shopping in response to a SSB tax.³² The present study places these findings in a larger context. While media reports following the tax's implementation have sensationalized cross-boundary "soda trips," the magnitude of changes observed in this analysis suggest that SNAP participants may have migrated entire grocery shopping trips (rather than purchases of taxed goods alone) outside city boundaries. Furthermore, USDA estimates that, on average, SNAP households make 9.4 transactions per month, with an average transaction amount of \$27.36.¹⁰ Existing research on food shopping among low-income Philadelphians adds to the plausibility of entire trips being migrated by suggesting that retailer proximity, while important, is far from the only factor that influences where to shop, which can include broader economic, logistical, and social considerations.^{33,34}

With an excise tax, applied at the distributor level, important questions exist about how much of the tax is passed through to customers via product prices, if any, and on which products. One could imagine an excise tax directly

passed through via proportionately higher prices on taxed goods, or more indirectly spread among taxed and untaxed products. A recent study of tax implementation in Philadelphia supermarkets found the mean price per ounce of taxed beverages increased by .83 cents from 2016 to 2017, and sales of taxed beverages decreased in Philadelphia and increased in neighboring areas, suggesting that price pass-through did occur.² These higher prices, in turn, could have motivated price-sensitive shoppers to shop outside Philadelphia, especially in neighborhoods near the county border. SNAP is predominantly spent at supermarket retailers, typically with the largest shopping trip occurring at the beginning of the month as benefits are renewed, and involving some form of private vehicle use, even by those who are not vehicle owners, but who might share rides with friends or family.^{15,33,34} Thus, if the price differences resulting from the tax were enough to motivate SNAP shoppers to shift these larger, vehicle-oriented trips across city borders, a sizeable effect might be expected.

Looking More Closely at Retailer Behavior

While higher prices on SSBs may have motivated shopping outside the city among low-income individuals, more explanation is needed. For instance, what role did retailer pricing and promotion strategies (on both taxed and untaxed items) play? Evidence from New York State reveals significant increases in in-store marketing of SSBs during periods of SNAP benefit issuance compared with other times of the month, with no attendant increases in marketing of low-calorie or

unsweetened beverages,³⁵ illustrating how retailers can target SNAP shopping. Although additional research is needed to unpack the pricing and promotion dynamics that may have motivated SNAP shoppers to move entire shopping trips across city boundaries, it is likely that retailers signaled to consumers beyond pricing alone.

Public Health Implications

In light of these and other findings that suggest shoppers may have moved rather than reduced purchases of taxed beverages, alternative strategies for achieving public health benefits should be considered. At the consumer level, a direct consumer tax, such as sales taxes more commonly associated with alcohol or tobacco, could be set uniformly across city retailers, rather than the general increase in prices that result from distributor and retailer pass-through of an excise tax. Still, this does little to solve the problem of cross-border shopping to avoid the tax, and other research suggests that sales taxes are less salient to consumers than excise taxes.³⁶ Notably, purchases made with SNAP are exempt from sales taxes, so these efforts may do little to reduce SSB consumption among SNAP shoppers.

In terms of retailer-level alternatives, policies intended to reduce consumption of certain goods are unlikely to win allies from the business community. Complementary policies or programs that encourage substitution of goods (e.g., water or unsweetened beverages) or additional purchasing (e.g., "double-up" bonuses for fruit and vegetable purchases) are perhaps more palatable to retailers, and also create both push and pull dynamics to help encourage

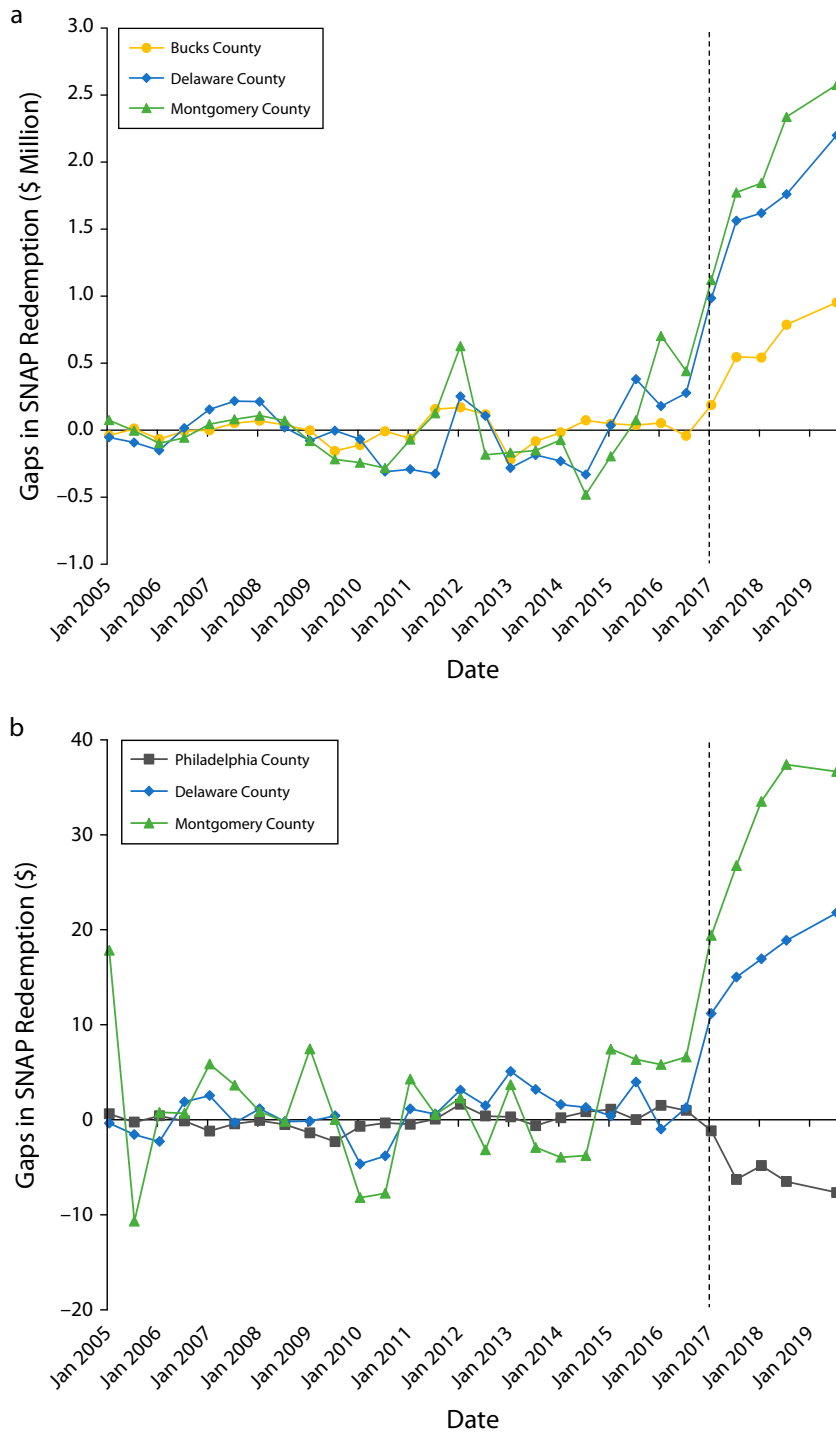


FIGURE 2— Gaps in Synthetic and Treated Counties for (a) Total SNAP Redemption and (b) SNAP Redemption per Participant: Philadelphia, PA, and Neighboring Pennsylvania Counties, 2005-2019

Note. SNAP = Supplemental Nutrition Assistance Program. Dashed vertical line indicates the implementation of the Philadelphia tax in January 2017. Gap values illustrate the estimated difference between treated and synthetic counties. SNAP redemption (panel A) synthetic controls were fit using log-transformed outcome values, which were back-transformed to dollar values here. No appropriate synthetic control could be fit for log SNAP redemption in Philadelphia County (panel A), and no detectable treatment effect was found for Bucks County (panel B), so they are not presented here.

healthy behavior changes at the population level.

Strengths and Limitations

The synthetic control approach offers a data-driven method that, when carefully applied, helps rule out alternative explanations by matching and weighting donor units based on underlying relationships between variables in a data set. This study merges several discrete county-level, time-series data sets related to the allocation and spending of SNAP benefits. Thus, we can more accurately consider the supply-and-demand dynamics related to SNAP spending.

This study also had several limitations. Here, it was assumed that households did not choose to stop participating in SNAP on the basis of the implementation of the beverage tax. It is also critical to note that this study did not examine items purchased by SNAP shoppers. While we can make comparisons to existing literature on purchasing patterns among SNAP participants, we cannot conclude that the price of SSBs specifically was the cause of the observed shift in SNAP spending, nor do we estimate individual-level changes in spending. More detailed consumer panel data may reveal if particular items motivated the broader shifts in spending.

Conclusions

The 2017 implementation of an SSB excise tax in Philadelphia coincided with nationwide decreases in SNAP enrollment. I used a synthetic control approach to estimate the average treatment effect of the tax on Philadelphia as well as its immediate neighbors and observed significant treatment

effects on SNAP shopping in Philadelphia's neighboring Pennsylvania counties, potentially indicating a shift in SNAP purchasing away from Philadelphia. However, the magnitude of these effects suggests that they are not likely to reflect SSB sales alone. More research is needed to understand the mechanisms involving both retailer and consumer behavior that might explain these responses to the excise tax. **AJPH**

ABOUT THE AUTHOR

Benjamin W. Chrisinger is with the Department of Social Policy and Intervention, University of Oxford, Oxford, United Kingdom.

CORRESPONDENCE

Correspondence should be sent to Benjamin W. Chrisinger, Barnett House, 32 Wellington Square, Oxford OX1 2ER, United Kingdom (e-mail: benjamin.chrisinger@spi.ox.ac.uk). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONFLICTS OF INTEREST

The author has no conflicts of interest to declare.

HUMAN PARTICIPANT PROTECTION

No human participants were involved in this analysis.

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Countermarketing About Fruit Drinks, Alone or With Water Promotion: A 2019 Randomized Controlled Trial in Latinx Parents

James Krieger, MD, MPH, Taehoon Kwon, PhD, Rudy Ruiz, MPP, Lina Pinero Walkinshaw, MPH, Jiali Yan, MS, and Christina A. Roberto, PhD

 See also Dorfman, p. 1905.

Objectives. To test whether fruit drink countermarketing messages alone or combined with water promotion messages reduce Latinx parents' purchases of fruit drinks for children aged 0 to 5 years.

Methods. We performed a 3-arm randomized controlled online trial enrolling 1628 Latinx parents in the United States during October and November 2019. We assessed the effect of culturally tailored fruit drink countermarketing messages (fruit drink-only group), countermarketing and water promotion messages combined (combination group), or car-seat safety messages (control) delivered via Facebook groups for 6 weeks on parental beverage choices from a simulated online store.

Results. The proportion of parents choosing fruit drinks decreased by 13.7 percentage points in the fruit drink-only group (95% confidence interval [CI] = -20.0, -7.4; $P < .001$) and by 19.2 percentage points in the combination group (95% CI = -25.0, -13.4; $P < .001$) relative to control. Water selection increased in both groups.

Conclusions. Fruit drink countermarketing messages, alone or combined with water promotion messages, significantly decreased parental selection of fruit drinks and increased water selection for their children.

Public Health Implications. Countermarketing social media messages may be an effective and low-cost intervention for reducing parents' fruit drink purchases for their children. (*Am J Public Health*. 2021;111(11):1997-2007. <https://doi.org/10.2105/AJPH.2021.306488>)

Consumption of sugar-sweetened beverages (SSBs) is associated with adverse health outcomes among children that are inequitably distributed by race and income.¹⁻³ In the United States, nearly half of children aged 2 to 4 years consume an SSB on a given day,⁴ and SSB intake is highest among Latinx and Black children.^{5,6} Latinx children are the largest racial/ethnic group among children of color in the United States.⁷ Fruit drinks (fruit-flavored beverages containing added sugar) are the

most-consumed SSB among young children, including Latinx children.^{4,5,8}

Most of the many public health awareness campaigns aimed at decreasing SSB consumption have used messages about the sugar content and health effects of these beverages.⁹⁻¹⁵ However, misleading marketing may lead parents to believe fruit drinks are healthy beverages, contributing to high consumption.^{16,17} Countermarketing campaigns highlighting industry's misleading messages may encourage

healthier beverage choices and could complement the more traditional messages.¹⁸ Tobacco countermarketing has increased antitobacco attitudes and lowered smoking rates and may serve as a model for applying this approach to unhealthy foods and beverages.^{19,20}

Public health SSB campaigns have primarily used mass media channels for message delivery (Lina Pinero Walkinshaw, e-mail communication, July 15, 2019). Using social media to target messages may be a more cost-effective

and scalable approach.²¹ Consumers, including Latinx people, are increasingly using social media as a source of health information.^{22,23}

Despite the promise of countermarketing campaigns and social media as public health communication tools, few studies have evaluated their effects on parents' beverage choices for their children and children's SSB consumption. To address this gap and the need to reduce SSB consumption among Latinx children, we conducted a randomized controlled trial to test the hypothesis that fruit drink countermarketing messages delivered via Facebook groups would reduce the proportion of parents choosing fruit drinks for their children and reduce child fruit drink intake. We also assessed whether adding positive messages promoting the health benefits of water would enhance the effects of the negative countermarketing messages.

METHODS

The study was a parallel group, prospective, 3-arm, randomized controlled trial conducted in the United States between October 11 and November 22, 2019. We used Facebook groups to deliver fruit drink countermarketing messages to one intervention arm (fruit drink-only group), a combination of fruit drink countermarketing and water promotion messages to a second (combination intervention group), and car seat safety messages to an attention control arm.²⁴ The complete Standard Protocol Items: Recommendations for Interventional Trials (SPIRIT) protocol is available upon request.

Participants

A survey research firm (Galloway Research) recruited study participants

using its nationwide proprietary database (populated with respondents to randomized phone survey invitations and former research project participants) and targeted social media messages. Potential participants were asked to enroll in a research study and share opinions about beverages. Eligibility criteria were self-identifying as Latinx, age 18 years or older, caretaker of a child aged 0 to 5 years, daily social media use, and preferring either English or Spanish when speaking. We stratified recruitment so that 80% of participants preferred English when using social media and 20% preferred Spanish, and educational status and household income were representative of the US Latinx population.²⁵

Study Procedures and Randomization

After online eligibility screening, eligible respondents were directed to an online baseline survey and offered \$15 for completing it. Before starting the survey, participants received information about study procedures and indicated that they understood and agreed to participate. We informed them that they would be asked to join a Facebook group for 6 weeks focused on kids' drinks and would receive information about these drinks. Those completing the survey were enrolled in the study and randomized to a study arm using a computer-generated random number and blocking procedure with randomly varied block size (Voxco RAN[0.01,0.99] command²⁶). All research team members were blinded to assignments until data collection was complete.

Within study arms, participants were assigned to an English- or Spanish-language Facebook group based on their language preference.

Each of 9 campaign messages was posted twice to the Facebook groups, so participants received 3 messages per week over a 6-week study period. Messages consisted of a short text header and image (Figure A, available as a supplement to the online version of this article at <http://www.ajph.org>). Participants were asked to set Facebook notifications to display messages in their News Feeds and to view the group page at least weekly and received weekly e-mail reminders to view messages. They could "like" and comment on the messages. Interlex, a Latinx-led bilingual communications team, monitored posts for compliance with group rules, but otherwise study team members did not interact with the participants in the Facebook groups.

After 6 weeks, all randomized participants received a link to an online exit survey, along with multiple e-mails and Facebook reminders to complete it. Participants received \$20 for completing the exit survey.

Intervention Development

We developed preliminary messages informed by a Centers for Disease Control and Prevention compilation of existing SSB message campaigns,⁹ a literature review, an Internet search, and an expert advisory group. Interlex created initial messages modeled on branding of fruit drinks popular in Latinx communities. Messages were developed simultaneously in English and Spanish to ensure consistency and relevance across languages and cultures. Findings from 5 focus groups across the United States with 45 Latinx parents of children aged 0 to 5 years informed the final content of the messages (Figure B, available as a

supplement to the online version of the article at <http://www.ajph.org>). For example, one message consists of an image of a child with tooth decay along with a fruit drink pouch displaying an “all-natural” claim and the text, “Just because a label states ‘all-natural’ doesn’t make a fruit drink healthy. Don’t let the beverage industry harm your kids.”

Measures

Full measures and surveys are in Appendices A and B (available as supplements to the online version of this article at <http://www.ajph.org>). The baseline survey included the following:

- Supermarket shopping task: We asked participants to imagine a typical trip to the supermarket or grocery store and buying a beverage for their oldest child between age 0 and 5 years. They were asked to select 1 beverage from an image of shelves containing 2 waters (both with no added sweeteners, vitamins, or additives, and 1 with and 1 without flavor), 1 soda, two 100% juices, 6 fruit drinks, and 1 milk (2% fat, unflavored; Figure C, available as a supplement to the online version of this article at <http://www.ajph.org>). To incentivize real-world shopping behavior, participants were told that once they completed the study, they would receive two \$2 coupons for the drinks they selected in the study store that could be redeemed at a real-world store (in reality they received a \$4 cash payment in lieu of the 2 promised \$2 beverage coupons). We assessed the proportion of parents choosing a fruit drink (primary outcome), water, soda, 100% juice, and milk.
 - Beverage perceptions: Participants viewed 4 beverage images (2 fruit drinks, 1 water, 1 soda) and answered questions about each one: “How much do you think your child would **enjoy** this drink?” and “How likely are you to **serve or buy** this drink for your child in the next four weeks?” from not at all to extremely. They then rated how strongly they agreed or disagreed with the following: “Drinking this product often would . . . (1) lead my child to **gain weight**; (2) increase my child’s risk of **diabetes**; (3) increase my child’s risk of **cavities and tooth problems**; (4) help my child live a **healthier** life.” All responses were on 1-to-7 Likert scales. We computed a health risk index by averaging responses to these last 4 items (reverse-coding the last statement); lower scores indicated healthier beverage perceptions. Fruit drink scores were averaged across the 2 drinks.
 - Beverage intake: We assessed children’s beverage consumption (ounces per day, main secondary outcome) with the Children’s BevQ beverage frequency questionnaire²⁷ and adult consumption (frequency per day) with 2 questions from the Family Life, Activity, Sun, Health, and Eating (FLASHE) Study.²⁸
 - Sociodemographic information: Participants provided their age, gender, race, country of origin, educational attainment, child’s age and gender, and household composition and income (from which we calculated percentage of the 2020 federal poverty level).²⁹
- The exit survey included the following additional measures:
- Message recall: Participants viewed each of the 9 messages they received and indicated how often they saw each in the past 6 weeks: never, once, or more than once. Recall was averaged across the 9 messages.
 - Message perceptions: Participants viewed 3 randomly selected messages from their study arm and provided Likert scale ratings on 10 dimensions including likeability, believability, and providing new information.^{30,31}
 - Perceptions of beverage brands: Participants viewed images of 2 fruit drink brands and completed a Net Promoter Score rating and an adaptation of the Brand Trust Scale.³² The Net Promoter Score assesses the likelihood of recommending a brand to a friend or colleague.³³ Based on response to a Likert scale ranging from 0 (not at all likely to recommend) to 10 (extremely likely to recommend), the respondent was classified as a promoter (9 or 10), passive (7 or 8), or detractor (0–6). The net score is the difference in the percentage of respondents who are promoters and detractors. We averaged the 5 items from the Brand Trust Scale to create an overall score from 1 to 5, with higher scores indicating greater trust.
 - Facebook engagement metrics: We collected data on the counts of “views,” “likes,” and comments for each message 1 week after posting.

Sample Size

A sample size of 385 per group was needed to detect a reduction of 10 percentage points in choice of fruit drink relative to control (conservatively assuming baseline prevalence of 50%), with 2-sided α of .05 and power of .8.

We sought to enroll 514 people per group to allow for 25% attrition.

Statistical Methods

The primary analysis was intention to treat and included all randomized participants regardless of whether they joined a Facebook group or completed the exit survey. We used multiple imputation with predictive mean matching using 20 imputation data sets, 84 variables, and a set of 5 candidate donors from complete cases for the missing entry to estimate missing exit survey outcome data for those not completing the study and for missing baseline data for parental beverage consumption (22% of latter missing because of a survey programming error).^{34,35} A secondary per-protocol analysis included only those participants who both joined a Facebook group and completed the exit survey.

We used linear regression models with robust standard errors to model the effect size of the intervention as the adjusted absolute difference in proportions (percentage point difference) or means between groups. The dependent variable was the postintervention value of the outcome. Independent variables were the baseline value of the outcome (except when outcome was measured at exit only) and a dummy variable for each intervention group (control was the reference).

We corrected model-generated P values for multiple comparisons.³⁶ We considered a corrected P value of less than .05 significant.

We conducted exploratory analyses to assess whether a set of demographic and baseline consumption variables modified intervention effects on the primary outcome with separate regression models for each variable. Models included the primary model variables

and terms that interacted the modifier with the group assignment variable. Because these were exploratory analyses, we deemed $P < .05$ without correction for multiple comparisons as significant. All analyses were 2-tailed and performed using R version 3.6.3³⁷ (including the {mice} package for multiple imputation and {miceadd} for linear model cluster robust standard error) and STATA/SE version 15.1.³⁸ We pre-registered the trial before recruitment at AsPredicted #29421.³⁹

RESULTS

We assessed 5297 individuals for eligibility and enrolled 1628 into the study, of whom 90% joined a Facebook group, and 79% completed both the intervention and the exit survey (Figure 1). Nine participants left the groups. Study arms were well-balanced demographically at baseline (Table 1). Participants were predominantly female and of Mexican descent, used social media frequently, lived in lower-income households, and had not completed college.

The proportions completing the study were similar across the groups (77%–82%). Participant characteristics associated with not joining a Facebook group included Spanish as preferred language, non-White race, less than college educational attainment, and low household income (Table A, available as a supplement to the online version of this article at <http://www.ajph.org>).

Primary Outcome

The proportion of parents choosing a fruit drink for their child decreased absolutely by 13.7 percentage points in the fruit drink–only group (95% confidence interval [CI] = $-20.0, -7.4$; $P < .001$) and by 19.2 percentage

points in the combination group (95% CI = $-25.0, -13.4$; $P < .001$) relative to the control group (Table 2). The decreases in the intervention groups did not differ significantly. The relative percent decrease compared with the control group was 30.9% (95% CI = 16.7%, 45.1%) for the fruit drink–only group and 42.6% (95% CI = 29.7%, 55.4%) for the combination group. The per-protocol analysis showed larger effect sizes and a significantly larger decrease in the combination arm relative to the fruit drink–only arm (Table B, available as a supplement to the online version of this article at <http://www.ajph.org>).

Secondary Outcomes and Additional Analyses

Choice of other beverages. The proportion of parents choosing water for their child (Table 2) increased by 17.6 percentage points in the fruit drink–only group relative to the control group (95% CI = 11.9, 23.4; $P < .001$) and by 29.7 percentage points in the combination group (95% CI = 24.0, 35.5; $P < .001$). The increase was larger in the combination group compared with the fruit drink–only group (12.1 percentage points; 95% CI = 5.7, 18.5; $P = .002$). The proportion choosing 100% fruit juice declined significantly in the combination group relative to control, although the difference between intervention groups was not significant. There were no significant differences in changes in choice of soda or milk across groups. Per-protocol analysis (Table B) showed a similar pattern but with larger effect sizes on choice of water and 100% juice.

Beverage intake. Children's fruit drink consumption decreased 0.6 ounces per

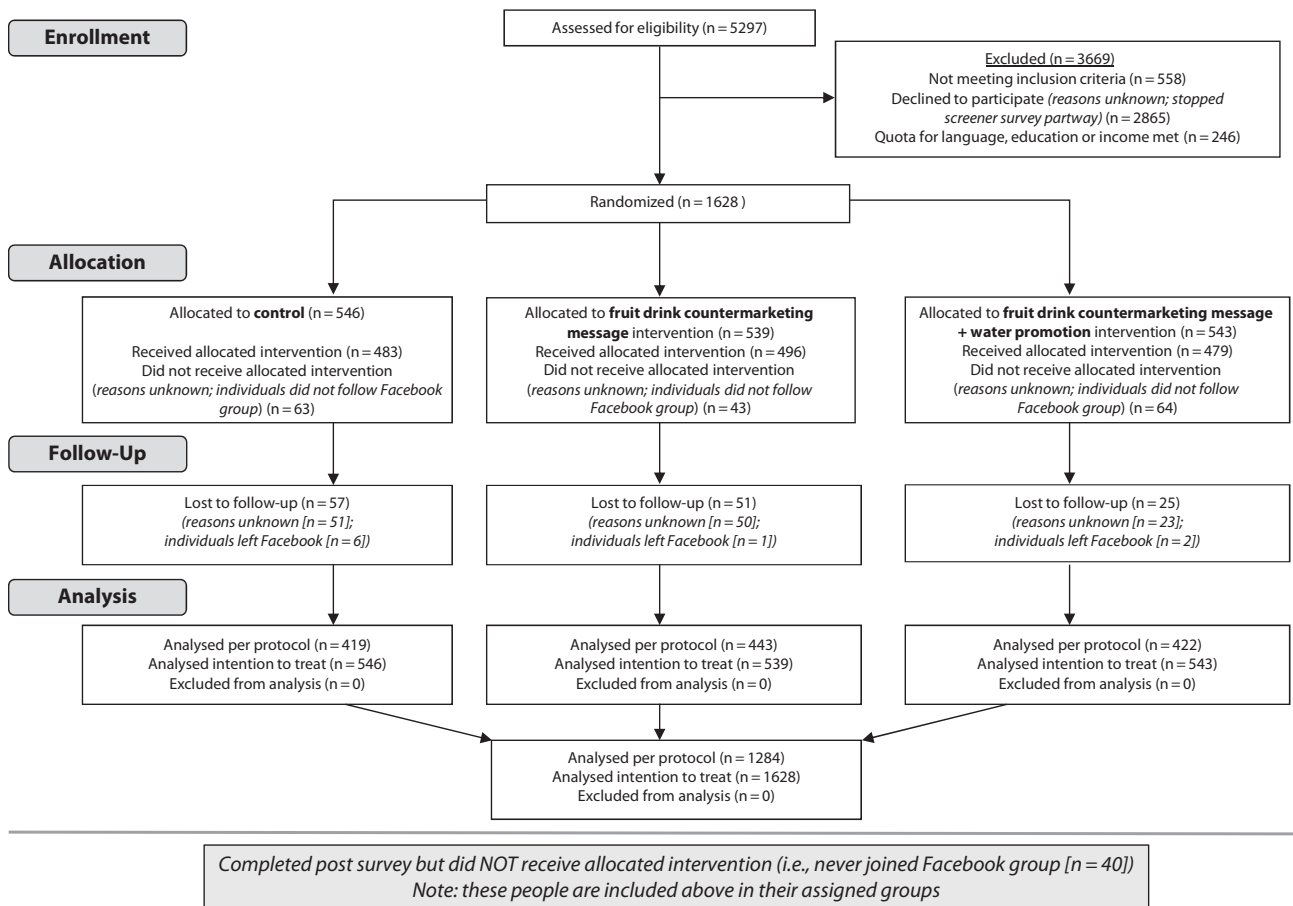


FIGURE 1— CONSORT Flow Diagram for Randomized Controlled Trial Participants

Note. CONSORT = Consolidated Standards of Reporting Trials.

day in the fruit drink-only group (95% CI = -1.1, -0.2; $P = .02$) and 0.8 ounces per day in the combined group (95% CI = -1.4, -0.3; $P = .01$) compared with the control group (Table 2). A decrease of 0.8 ounces per day is equivalent to approximately 2 grams of added sugars, given the added sugars content of commonly consumed fruit drinks. This represents a 22% decrease in added sugars from sweetened beverages consumed by children aged 2 to 5 years.⁴⁰ Consumption of milk and 100% fruit juice decreased significantly more in both intervention groups relative to the control group. We observed no significant differences in changes in water

consumption. Parents in both intervention groups reported drinking sugary drinks less frequently compared with the control group. The per-protocol analysis revealed similar patterns for parents' and children's beverage consumption, although children's reductions for all types of SSB intake were significant for both arms (data not shown).

Beverage perceptions. Parental perceptions of fruit drinks as healthy or enjoyable and their intentions to buy or serve them for their children decreased significantly in both intervention groups relative to the control group (Table 3). Parents' intention to serve or buy soda

for their children decreased significantly in both groups while perceptions of soda as healthy and enjoyable declined only in the combined group.

Brand perceptions and trust. Parents in both intervention groups were significantly less likely to be promoters of fruit drink brands and to trust fruit drink brands at the time of the exit survey relative to the control group (Table 3).

Tests for effect modification. We did not detect significant modification of the intervention effect on the primary outcome by caretaker age, income,

TABLE 1— Baseline Characteristics of 1628 Participants in a Randomized Controlled Trial of a Social Media Fruit Drink Countermarketing Campaign for Latinx Parents of Children Aged 0 to 5 Years: United States, 2019

Characteristic	Study Group, Mean \pm SD or No. (%)			
	Overall (n = 1628)	Fruit Drink Only (n = 539)	Combination (n = 543)	Control (n = 546)
Parent				
Age, y ^a	30.94 \pm 5.16	30.76 \pm 5.13	31.03 \pm 5.26	31.03 \pm 5.09
Gender ^b				
Female	1483 (91.1)	482 (89.4)	498 (91.7)	503 (92.1)
Male	140 (8.6)	55 (10.2)	43 (7.9)	42 (7.7)
White race	828 (50.9)	277 (51.4)	280 (51.6)	271 (49.6)
Mexican origin or descent	1129 (69.3)	374 (69.4)	385 (70.9)	370 (67.8)
Educational attainment ^b				
High school or less	585 (35.9)	199 (36.9)	183 (33.7)	203 (37.2)
Some college attended, but no degree	526 (32.3)	182 (33.8)	172 (31.7)	172 (31.5)
College degree or higher	493 (30.3)	150 (27.8)	179 (33.0)	164 (30.0)
Preferred language when speaking with others				
English more than Spanish	718 (44.1)	253 (46.9)	232 (42.7)	233 (42.7)
English and Spanish equally	595 (36.5)	189 (35.1)	201 (37)	205 (37.5)
Spanish more than English	315 (19.3)	97 (18.0)	110 (20.3)	108 (19.8)
Prefer to use English when using social media	1275 (78.3)	432 (80.1)	411 (75.7)	432 (79.1)
Use Facebook more than once a day	1512 (92.9)	497 (92.2)	508 (93.6)	507 (92.9)
Use Instagram more than once a day	1066 (65.5)	368 (68.3)	340 (62.6)	358 (65.6)
Child				
Age, y	3.32 \pm 1.42	3.34 \pm 1.41	3.34 \pm 1.44	3.30 \pm 1.41
Gender ^b				
Female	815 (50.1)	282 (52.3)	261 (48.1)	272 (49.8)
Male	802 (49.3)	253 (46.9)	278 (51.2)	271 (49.6)
Household				
Income as % of federal poverty level ²⁹				
< 100%	682 (41.9)	232 (43.0)	215 (39.6)	235 (43)
100% to 199%	505 (31.0)	174 (32.3)	170 (31.3)	161 (29.5)
200% to 399%	313 (19.2)	96 (17.8)	110 (20.3)	107 (19.6)
\geq 400%	128 (7.9)	37 (6.9)	48 (8.8)	43 (7.9)
No. of adults in household				
1	189 (11.6)	62 (11.5)	67 (12.3)	60 (11.0)
2	1116 (68.6)	367 (68.1)	363 (66.9)	386 (70.7)
\geq 3	323 (19.8)	110 (20.4)	113 (20.8)	100 (18.3)
No. of children in household				
1	449 (27.6)	126 (23.4)	154 (28.4)	169 (31.0)
2	561 (34.5)	192 (35.6)	199 (36.6)	170 (31.1)
3	372 (22.9)	130 (24.1)	117 (21.5)	125 (22.9)
\geq 4	246 (15.1)	91 (16.9)	73 (13.4)	82 (15.0)

^aAge is missing for 23 participants.

^bPercentages do not add to 100% because of unknown or refused responses.

TABLE 2— Beverage Choice and Consumption: Baseline and Exit Values by Study Group and Adjusted Differences Across Study Groups: United States, 2019

	Fruit Drink Only (n = 539), Mean (95% CI) or % (No.) ^a		Combination (Fruit Drink + Water) (n = 543), Mean (95% CI) or % (No.) ^a		Control (n = 546), Mean (95% CI) or % (No.) ^a		Fruit Drink Only Vs Control, Difference (95% CI) ^b	Combination vs Control, Difference (95% CI) ^b	Combination vs Fruit Drink Only, Difference (95% CI) ^b
	Baseline	Exit	Baseline	Exit	Baseline	Exit			
Parent beverage choice in online store, % choosing ^c									
Fruit drink	44.3 (239)	27.4 (148)	45.1 (245)	22.1 (120)	47.6 (260)	41.9 (229)	-13.7*** (-20.0, -7.4)	-19.2*** (-25.0, -13.4)	-5.5 (-11.6, 0.7)
Soda	1.5 (8)	2.1 (11)	2.0 (11)	2.2 (12)	2.0 (11)	3.8 (21.0)	-1.6 (-3.7, 0.6)	-1.6 (-3.7, 0.4)	-0.1 (-1.9, 1.8)
Water	22.6 (122)	40.7 (219)	21.5 (117)	52.5 (285)	20.0 (109)	22.3 (121)	17.6*** (11.9, 23.4)	29.7*** (24.0, 35.5)	12.1** (5.7, 18.5)
Milk	8.5 (46)	14.1 (76)	8.3 (45)	10.1 (55)	9.2 (50)	11.4 (62)	2.9 (-1.3, 7.0)	-1.1 (-5.0, 2.8)	-4.0 (-7.8, -0.1)
100% fruit juice	23.0 (124)	15.7 (85)	23.0 (125)	13.1 (71)	21.2 (116)	20.5 (112)	-5.2 (-10.0, -0.5)	-7.8* (-12.6, -3.1)	-2.6 (-7.1, 1.9)
Child beverage consumption, mean oz per day ^d									
Fruit drink	2.7 (2.2, 3.1)	1.7 (1.3, 2.0)	2.5 (2.1, 2.9)	1.4 (1.0, 1.9)	3.0 (2.6, 3.5)	2.4 (2.0, 2.8)	-0.6* (-1.1, -0.2)	-0.8* (-1.4, -0.3)	-0.2 (-0.7, 0.4)
Regular soda	1.1 (0.8, 1.4)	0.7 (0.5, 0.9)	1.2 (0.9, 1.5)	0.8 (0.4, 1.1)	1.2 (1.0, 1.5)	1.2 (0.8, 1.5)	-0.4 (-0.7, -0.1)	-0.4 (-0.8, 0.1)	0.0 (-0.3, 0.4)
Diet soda	0.4 (0.2, 0.6)	0.4 (0.1, 0.6)	0.7 (0.4, 0.9)	0.3 (0.1, 0.6)	0.6 (0.4, 0.8)	0.7 (0.4, 1.0)	-0.2 (-0.6, 0.2)	-0.4 (-0.8, 0.0)	-0.1 (-0.5, 0.2)
Sports and energy drink	1.0 (0.8, 1.3)	0.7 (0.5, 0.9)	1.1 (0.8, 1.4)	0.6 (0.2, 1.1)	1.3 (1.0, 1.6)	1.2 (0.8, 1.5)	-0.4 (-0.7, 0.0)	-0.4 (-1.0, 0.1)	-0.1 (-0.6, 0.5)
Sweet tea	0.7 (0.5, 0.9)	0.5 (0.3, 0.7)	0.7 (0.4, 0.9)	0.4 (0.1, 0.8)	0.8 (0.5, 1.1)	0.8 (0.5, 1.2)	-0.3 (-0.6, 0.0)	-0.4 (-0.8, 0.1)	0.0 (-0.4, 0.3)
Water	15.8 (15.1, 16.6)	15.1 (14.2, 15.9)	16.3 (15.5, 17.1)	16.7 (15.9, 17.6)	16.1 (15.3, 16.9)	15.6 (14.8, 16.5)	-0.4 (-1.4, 0.6)	1.0 (0.0, 2.0)	1.4 (0.4, 2.4)
Milk	10.1 (9.4, 10.9)	8.1 (7.3, 8.8)	10.1 (9.4, 10.8)	7.8 (7.1, 8.5)	9.8 (9.1, 10.5)	9.3 (8.5, 10.1)	-1.4* (-2.3, -0.6)	-1.7* (-2.6, -0.8)	-0.3 (-1.1, 0.6)
100% fruit juice	4.2 (3.8, 4.7)	2.7 (2.2, 3.2)	4.7 (4.2, 5.2)	2.6 (1.9, 3.2)	4.7 (4.1, 5.2)	4.1 (3.5, 4.7)	-1.2** (-1.9, -0.6)	-1.5*** (-2.3, -0.7)	-0.3 (-1.1, 0.4)
Other ^e	0.1 (0.1, 0.2)	0.3 (0.1, 0.5)	0.3 (0.2, 0.5)	0.1 (0.0, 0.3)	0.3 (0.1, 0.5)	0.2 (0.1, 0.4)	0.1 (-0.2, 0.3)	-0.1 (-0.3, 0.1)	-0.2 (-0.4, 0.1)
All sugar sweetened beverages	5.6 (4.8, 6.4)	3.6 (2.9, 4.3)	5.5 (4.5, 6.4)	3.3 (2.0, 4.6)	6.4 (5.4, 7.3)	5.6 (4.4, 6.7)	-1.7 (-2.8, -0.6)	-1.9 (-3.5, -0.3)	-0.2 (-1.7, 1.2)
All healthier drinks ^f	26.0 (24.8, 27.2)	23.1 (22.0, 24.3)	26.4 (25.3, 27.5)	24.5 (23.4, 25.6)	25.9 (24.8, 27.0)	25.0 (23.7, 26.2)	-1.8 (-3.2, -0.4)	-0.7 (-2.1, 0.8)	1.2 (-0.3, 2.6)
Parent beverage consumption, mean times per day ^g									
Fruit drinks and teas	0.4 (0.4, 0.5)	0.3 (0.3, 0.4)	0.5 (0.4, 0.5)	0.3 (0.3, 0.4)	0.5 (0.4, 0.5)	0.4 (0.3, 0.4)	0.0 (-0.1, 0.0)	0.0 (-0.1, 0.0)	0.0 (-0.1, 0.1)
All sugary drinks	0.9 (0.8, 0.9)	0.6 (0.6, 0.7)	0.8 (0.7, 0.9)	0.6 (0.5, 0.6)	0.9 (0.8, 0.9)	0.7 (0.7, 0.8)	-0.1* (-0.2, 0)	-0.2* (-0.3, -0.1)	0.0 (-0.1, 0.1)

Note. CI = confidence interval.

^aBaseline and exit values for proportions within groups were modeled using linear regression with robust standard error, which gives the percent value (%). No. is calculated using (n) × (total n).

^bAdjusted percentage point differences between groups were modeled using linear regression with robust standard error controlling for baseline value. Adjusted mean differences between groups were modeled using linear regression with robust standard error controlling for baseline value.

^cP values of between-group comparisons were adjusted using Bonferroni-Holm correction for 3 tests for fruit drinks (primary prespecified outcome) and 12 tests for other beverages.

^dP values of between-group comparisons were adjusted using Bonferroni-Holm correction for 3 tests for fruit drinks (secondary prespecified outcome) and 30 tests for other beverages.

^eIncludes aguas frescas, nut milk, tea, coffee, smoothies, hot chocolate, yogurt drinks, and multiple beverages combined in self-report.

^fIncludes milk and water.

^gP values of between-group comparisons were adjusted using Bonferroni-Holm correction for 6 tests.

*P < .05; **P < .01; ***P < .001.

TABLE 3— Beverage Health Perceptions and Brand Perceptions: Baseline and Exit Values by Study Group and Adjusted Differences Across Study Groups: United States, 2019

	Fruit Drink Only (n = 539), Mean (95% CI) or % (No.)		Combination (Fruit Drink + Water; n = 543), Mean (95% CI) or % (No.)		Control (n = 546), Mean (95% CI) or % (No.)		Fruit Drink-Only Vs Control, Difference (95% CI)	Combination Vs Control, Difference (95% CI)	Combination Vs Fruit Drink Only, Difference (95% CI)
	Baseline	Exit	Baseline	Exit	Baseline	Exit			
Parent beverage perceptions, mean rating^{a,b}									
Child's enjoyment of beverage ^c									
Fruit drink	4.9 (4.7, 5.0)	4.5 (4.3, 4.7)	4.8 (4.6, 4.9)	4.3 (4.1, 4.5)	4.8 (4.6, 4.9)	4.8 (4.6, 5.0)	-0.4** (-0.6, -0.2)	-0.5*** (-0.7, -0.3)	-0.1 (-0.3, 0.1)
Soda	4.2 (4.0, 4.4)	4.2 (4.0, 4.4)	4.0 (3.8, 4.2)	3.8 (3.5, 4.0)	4.2 (4.0, 4.4)	4.5 (4.3, 4.7)	-0.3 (-0.5, 0.0)	-0.6*** (-0.8, -0.3)	-0.3 (-0.5, 0.0)
Water	5.8 (5.6, 5.9)	5.9 (5.7, 6.0)	5.8 (5.7, 5.9)	6.0 (5.6, 6.4)	5.7 (5.6, 5.8)	5.8 (5.6, 6.0)	0.0 (-0.2, 0.2)	0.2 (-0.3, 0.6)	0.1 (-0.3, 0.6)
Likelihood of serving or buying for child ^c									
Fruit drink	3.2 (3.1, 3.3)	2.1 (2.0, 2.3)	3.3 (3.1, 3.4)	2.2 (2.1, 2.3)	3.1 (3.0, 3.2)	3.1 (2.9, 3.2)	-1.0*** (-1.1, -0.8)	-0.9*** (-1.1, -0.8)	0.0 (-0.1, 0.2)
Soda	2.1 (2.0, 2.2)	1.8 (1.7, 1.9)	2.2 (2.0, 2.3)	1.8 (1.7, 1.9)	2.2 (2.0, 2.3)	2.3 (2.1, 2.4)	-0.4*** (-0.6, -0.3)	-0.5*** (-0.6, -0.3)	-0.1 (-0.2, 0.1)
Water	6.4 (6.3, 6.5)	6.6 (6.5, 6.7)	6.3 (6.2, 6.5)	6.6 (6.4, 6.7)	6.3 (6.2, 6.4)	6.5 (6.3, 6.6)	0.1 (0.0, 0.3)	0.1 (-0.1, 0.3)	0.0 (-0.2, 0.1)
Health risk of beverage to child ^d									
Fruit drink	4.6 (4.5, 4.7)	5.6 (5.5, 5.7)	4.6 (4.5, 4.7)	5.5 (5.3, 5.6)	4.6 (4.5, 4.7)	4.7 (4.6, 4.8)	0.9*** (0.7, 1.0)	0.7*** (0.6, 0.9)	-0.1 (-0.3, 0.0)
Soda	6.4 (6.3, 6.5)	6.5 (6.4, 6.6)	6.4 (6.3, 6.5)	6.6 (6.5, 6.6)	6.4 (6.3, 6.4)	6.4 (6.3, 6.4)	0.1 (0.0, 0.2)	0.2* (0.1, 0.3)	0.0 (-0.1, 0.2)
Water	1.4 (1.4, 1.5)	1.4 (1.3, 1.5)	1.4 (1.3, 1.5)	1.3 (1.2, 1.4)	1.4 (1.3, 1.5)	1.4 (1.3, 1.6)	-0.1 (-0.2, 0.1)	-0.2 (-0.3, 0.0)	-0.1 (-0.2, 0.0)
Parent brand perceptions at exit, % in each category (n)^{e,f}									
Promoters	...	4.5 (24)	...	5.6 (30)	...	17.3 (94)	-12.8*** (-16.8, -8.7)	-11.7*** (-15.9, -7.5)	1.1 (-1.8, 4.0)
Passives	...	7.1 (38)	...	9.9 (54)	...	24.4 (133)	-17.3*** (-21.7, -12.9)	-14.5*** (-19.2, -9.9)	2.8 (-0.8, 6.4)
Detractors	...	88.4 (477)	...	84.5 (459)	...	58.3 (318)	30.1*** (25.0, 35.3)	26.2*** (20.6, 31.9)	-3.9 (-8.3, 0.5)
Parent brand trust at exit, mean score (95% CI)^{g,h,i}									
Brand trust score	...	2.3 (2.1, 2.4)	...	2.4 (2.3, 2.5)	...	3.4 (3.3, 3.5)	-1.1*** (-1.3, -1.0)	-0.9*** (-1.1, -0.8)	0.2* (0.0, 0.3)

Note. CI = confidence interval.

^aP values of between-group comparisons were adjusted using Bonferroni-Holm correction for 9 tests for each of the 3 measures.

^bAdjusted mean differences between groups were modeled using linear regression with robust standard error controlling for baseline value.

^cResponse from 1 to 7, where 1 means not at all, and 7 means extremely.

^dResponse from 1 to 7, where 1 means healthy, and 7 means unhealthy.

^eResponse options were a Likert scale from 0 (not at all likely to recommend brand to friend or colleague) to 10 (extremely likely). Participants were assigned to the following categories based on their scores: as promoter (9 or 10), passive (7 or 8), or detractor (0-6). The net promoter score was calculated as (percent promoters)-(percent detractors).

^fAdjusted percentages point differences between groups were modeled using linear regression with robust standard error.

^gP values of between-group comparisons were adjusted using Bonferroni-Holm correction for 3 tests.

^hResponse from 1 to 5, where 1 is completely disagree and 5 is completely agree with statements describing level of trust in brand (lower score means lower trust).

ⁱMean differences between groups were modeled using linear regression with robust standard error.

*P < .05; **P < .01; ***P < .001.

education, race, and preferred language, nor child's or caretaker's baseline fruit drink consumption (Table C, available as a supplement to the online version of this article at <http://www.ajph.org>).

Facebook Engagement, Messages, and Cost

The proportion of participants joining their assigned Facebook group ranged from 88.5% to 92.0% across groups. Averaged across all 6 weeks, 80.0% to 83.5% of group members viewed the Facebook posts. The proportion of participants "liking" messages from each group, averaged across all messages received during the 6 weeks, ranged from 36.4% to 40.5%. The average proportion posting comments ranged from 0% to 10.7%. One third (33.7%) of intervention-group participants reported seeing the messages at least once, while 47.1% reported seeing them more than once, and 19.2% reported never seeing them or did not know. Participants found the messages convincing and informative (Table D, available as a supplement to the online version of this article at <http://www.ajph.org>). The cost of delivering the messages and moderating the Facebook group was \$20 000, or \$12.29 per person.

DISCUSSION

The delivery of culturally tailored countermarketing messages about fruit drinks via a Facebook group to Latinx parents of young children, alone or combined with water promotion messages, led to large and significant reductions in the proportion of parents choosing fruit drinks for their children in an online simulated store. There was no effect

modification by race, income, education level, age, or language preference. The fruit drink messages also increased parents' selection of water, with a larger effect from the combined messages. Both messages led to significant decreases in parents' reports of their children's fruit drink consumption and increased parents' negative perceptions of fruit drinks and beverage company brands. To our knowledge, this is the first study to demonstrate the efficacy of countermessages delivered solely via social media as well as the first to specifically target sweetened beverage consumption among young children.

Unexpectedly, we observed a decrease in parental report of their children's consumption of both 100% juice and milk. Juice consumption may have declined because parents may not always clearly differentiate between 100% juice and fruit drinks containing some juice and may perceive the large total sugar content of 100% juice to be unhealthy. Milk consumption may have declined if parental concerns about artificial flavors in flavored milks increased and they therefore chose milk less often for their children.

Although SSB or added sugars countermarketing campaigns have been described, none have been rigorously evaluated.^{41,42} There are numerous reports of mass media campaigns using traditional health education messages—not countermessages—focused on the amount of added sugars in SSBs and their health effects.^{9–15} These campaigns have used multiple communication channels, including social media digital advertisements, but not social media groups. Evaluations have yielded mixed findings, and most have been limited to a single site or used an uncontrolled study design

(only 3 were controlled^{10,11,15}). The positive studies found 3% to 10% relative reductions in SSB sales or self-reported consumption. Costs ranged from \$300 000 to \$1.6 million in the 3 studies reporting them, which may be prohibitive for public and nonprofit agencies.^{10,11,13}

Many public health organizations use social media to disseminate health messages, but few rigorous studies have evaluated their impact. Most social media-based nutrition interventions have been tested in small pilot or feasibility studies and have focused primarily on White youths and adults.⁴³ Our study suggests that organic social media may be an effective, low-cost method for organizations with existing social media followings to launch SSB communication campaigns. As costs were fixed, the cost per person would decrease proportionate to the number of participants.

Our study had several limitations. First, our primary outcome was beverage choice in a simulated online store rather than a real store, but participants believed that they would receive a coupon for their beverage selection, incentivizing a real-world choice. Self-reported beverage consumption may be biased by social desirability, although we used a well-validated questionnaire and did not reveal study hypotheses to participants. Future research should assess additional outcomes like retail sales or 24-hour dietary recalls.

Second, we do not know whether effects persisted after the interventions ended. Third, participants were recruited in part from a marketing database and may not be representative of the population of Latinx parents of young children, although they do reflect

the educational status and household income of the US Latinx population. Fourth, people with lower incomes, with less education, who were non-White, and who preferred communicating in Spanish were less likely to join a Facebook group or complete the study. Barriers to engaging these populations in social media interventions and approaches for additional tailoring of interventions should be explored. Fifth, participants were recruited and offered incentives to join our social media groups, which might be prohibitively expensive in the real world. It will be useful to learn whether social media advertising alone, which is less costly than implementing social media groups, has effects comparable to those produced by joining a group. Finally, although effect modification analyses were not significant, our sample size was insufficient to detect modest effects.

Study strengths included its longitudinal, randomized controlled design; large sample size; high degree of participant engagement; and good participant retention.

Conclusions

In summary, this first, to our knowledge, rigorous assessment of fruit drink countermarketing messages tailored to Latinx parents of young children and delivered via a Facebook social media group, alone and in combination with water promotion messages, demonstrated reductions in parental choice of these beverages in an online simulated store and children's reported intake. Purchases of water increased. These results suggest that countermarketing messages delivered through social media groups may be a useful addition

to existing sugary drink–reduction strategies.

Public Health Implications

Our findings highlight the promise of social media countermarketing messages as either a low-cost stand-alone tool or one integrated into broader mass media campaigns, deployed during SSB policy adoption campaigns, or combined with other SSB reduction strategies.⁴⁴ It would be useful to understand the added benefits of combining social media countermarketing messages with traditional mass media campaigns and to test our approach with different beverage types and populations. **AJPH**

ABOUT THE AUTHORS

James Krieger and Lina Pinero Walkinshaw are with the Department of Health Systems and Population Health, School of Public Health, University of Washington, Seattle, WA. Taehoon Kwon is with the Department of Economics, University of Washington. Rudy Ruiz is with Interlex, San Antonio, TX. Jiali Yan and Christina A. Roberto are with the Department of Medical Ethics and Health Policy, Perelman School of Medicine, University of Pennsylvania, Philadelphia.

CORRESPONDENCE

Correspondence should be sent to James Krieger, MD, MPH, Department of Health Systems and Population Health, University of Washington, 3980 15th Ave N, Suite 415, Box 351621, Seattle, WA 98195 (e-mail: jwkrieg@uw.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

J. Krieger originated the study. J. Krieger and C. A. Roberto designed the study and provided oversight of all aspects of study implementation and analysis. L. P. Walkinshaw performed project management and focus group implementation

analysis. J. Krieger, L. P. Walkinshaw, and R. Ruiz designed the messages. J. Krieger, C. A. Roberto, L. P. Walkinshaw, and T. Kwon designed the survey and measures. J. Yan and T. Kwon analyzed the data. J. Krieger prepared drafts of the article with input from co-authors. All authors contributed to refinement of the study design and protocol and reviewed the final article.

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CONFLICTS OF INTEREST

The authors have no conflicts of interest to report.

HUMAN PARTICIPANT PROTECTION

Because the research team had no participant contact or access to identifiable data, the University of Washington institutional review board (IRB) determined the study did not require IRB review. The survey firm adhered to its established privacy policy (<https://www.gallowayresearch.com/privacy-policy>).

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Is the Probation Population's Health and Health Care Utilization Gendered?

Julia Dillavou, MA, and Yiwen Zhang, MA, MPP

ABOUT THE AUTHORS

The authors are PhD candidates in the Department of Sociology and Criminology, Pennsylvania State University, University Park.

See also Hawks et al., p. e2.

We read with great interest the article by Hawks et al.¹ comparing the health and care use of adults under probation with that of adults not under supervision. The authors used data from the National Survey on Drug Use and Health to draw comparisons between individuals 18 to 49 years of age who either were or were not on probation. We

commend the authors for focusing on the characteristics and unmet needs of this population; however, we have some concerns. Most importantly, given the evidence of gender differences in populations with psychiatric and substance use disorders^{2,3} as well as justice involvement,^{4,5} we wonder why the authors chose not to further split their analyses by gender or include interactions in their modeling strategy.

Figure 1 of the article displays the proportions of respondents on probation with physical and mental health conditions broken down by racial categories. Although salient in understanding the prevalence of such conditions and access to care, race is inextricably connected to gender within the power structures that affect social determinants of health and access to resources, including treatment. This figure shows little variation across race, but it does not address the evidence that justice-involved women have high rates of mental health diagnoses, substance use disorders, and histories of material hardship and trauma. These rates are often found to be higher than the rates among both justice-involved men and the general population of women.

It is imperative to consider gender when comparing justice-involved populations with the general population. Women entering the criminal justice system are likely to have a gendered set of risk factors, including a high prevalence of mental health problems, substance use disorder, and material hardship. Women also face unique barriers to entry for substance use disorder treatment and experience gender-based stigma for such disorders.^{6,7} These risk factors and barriers may be further exacerbated by racialized systems of disadvantage. We believe that by not accounting for gendered differences in condition prevalence and care use as well as potential interactions with race, Hawks et al. may have left significant associations out of their explanations for the distinctions between individuals on parole and those not under supervision. **AJPH**

CORRESPONDENCE

Correspondence should be sent to Julia Dillavou, MA, 211 Oswald Tower, University Park, PA 16802 (e-mail: julia.dillavou@psu.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

J. Dillavou conceptualized and contributed to the writing of the letter. Y. Zhang contributed to the writing and editing of the letter.

CONFLICTS OF INTEREST

The authors declare no potential conflicts of interest.

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Hawks et al. Respond

Laura Hawks, MD, MPH, Emily A. Wang, MD, MSc, Benjamin Howell, MD, MPH, Steffie Woolhandler, MD, MPH, David U. Himmelstein, MD, David Bor, MD, and Danny McCormick, MD, MPH

ABOUT THE AUTHORS

Laura Hawks is with the Medical College of Wisconsin, Milwaukee. Emily A. Wang and Benjamin Howell are with the Yale School of Medicine, New Haven, CT. Steffie Woolhandler and David U. Himmelstein are with Hunter College, City University of New York, New York, NY, and Harvard Medical School, Boston, MA. David Bor and Danny McCormick are with the Cambridge Health Alliance, Cambridge, MA, and Harvard Medical School.

 See also Dillavou and Zhang, p. e1.

We thank Dillavou and Zhang for their thoughtful comments on our study and for highlighting that gender-based barriers are often exacerbated by racialized systems of disadvantage.

We sought to characterize the health status of US adults on probation—a large and little studied population—with particular attention to Black and Native Americans, who have been disproportionately criminalized in the United States. Although space limitations precluded our exploring many potentially

important subgroups and interactions, we welcome the opportunity to address the questions raised by Dillavou and Zhang regarding possible interactions between gender and race/ethnicity.

We include here a gender-stratified version of Figure 1 from our article (Table 1), which showed the prevalence of physical and behavioral health conditions by race/ethnicity. As Dillavou and Zhang suspected, the prevalence of physical conditions and mental health disorders—but not substance use disorders—differs by gender. However, we find little evidence of

a sizeable or consistent interaction between gender and race.

We hope that our study will be a prelude to much additional research on the concomitants and effects of probation, including more detailed analyses of gender-based inequities. [AJPH](#)

CORRESPONDENCE

Correspondence should be sent to Laura Hawks, MD, MPH, Division of General Internal Medicine, 8701 Watertown Plank Rd, Milwaukee, WI 53226-3596 (e-mail: lhawks@mcw.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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CONTRIBUTORS

L. Hawks analyzed the data and drafted the response. L. Hawks and D. McCormick developed the analysis plan. E. A. Wang, B. Howell, S. Woolhandler, D. U. Himmelstein, and D. Bor made substantial contributions to the analysis plan and to the final version of the response.

CONFLICTS OF INTEREST

The authors report no conflicts of interest.

TABLE 1— Percentages of Adults on Probation (Stratified by Gender) With Physical Conditions, Mental Health Disorders, and Substance Use Disorders, by Race/Ethnicity: United States, 2015–2018

	More Than 1 Physical Condition		Serious Mental Illness		Any Mental Illness		Substance Use Disorder	
	Women (<i>P</i> = .01)	Men (<i>P</i> = .30)	Women (<i>P</i> = .05)	Men (<i>P</i> < .01)	Women (<i>P</i> < .01)	Men (<i>P</i> < .01)	Women (<i>P</i> = .02)	Men (<i>P</i> = .02)
White	42.6	28.3	22.4	11.6	53.1	32.4	37.3	37.2
Black	41.2	24.9	14.2	5.5	38.5	22.4	26.5	29.5
Hispanic	27.6	23.1	12.4	4.2	38.0	24.7	25.1	37.2
Native American	38.1	20.3	10.1	3.5	36.3	18.7	17.1	31.7

Note. Statistical significance levels are from the χ^2 test.

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Disparities in Health and Economic Burdens of Cancer Attributable to Suboptimal Diet in the United States, 2015–2018

Lu Wang, PhD, MPH, Mengxi Du, MS, MPH, RD, Frederick Cudhea, PhD, Christina Griecci, PhD, Dominique S. Michaud, ScD, Dariush Mozaffarian, MD, DrPH, and Fang Fang Zhang, MD, PhD

See also McCullough and Islami, p. 1913, and Galea and Vaughan, p. 1932.

Objectives. To quantify disparities in health and economic burdens of cancer attributable to suboptimal diet among US adults.

Methods. Using a probabilistic cohort state-transition model, we estimated the number of new cancer cases and cancer deaths, and economic costs of 15 diet-related cancers attributable to suboptimal intake of 7 dietary factors (a low intake of fruits, vegetables, dairy, and whole grains and a high intake of red and processed meats and sugar-sweetened beverages) among a closed cohort of US adults starting in 2017.

Results. Suboptimal diet was estimated to contribute to 3.04 (95% uncertainty interval [UI] = 2.88, 3.20) million new cancer cases, 1.74 (95% UI = 1.65, 1.84) million cancer deaths, and \$254 (95% UI = \$242, \$267) billion economic costs among US adults aged 20 years or older over a lifetime. Diet-attributable cancer burdens were higher among younger adults, men, non-Hispanic Blacks, and individuals with lower education and income attainments than other population subgroups. The largest disparities were for cancers attributable to high consumption of sugar-sweetened beverages and low consumption of whole grains.

Conclusions. Suboptimal diet contributes to substantial disparities in health and economic burdens of cancer among young adults, men, racial/ethnic minorities, and socioeconomically disadvantaged groups. (*Am J Public Health.* 2021;111(11):2008–2018. <https://doi.org/10.2105/AJPH.2021.306475>)

Cancer is a major public health burden and the second leading cause of death in the United States, with approximately 1.8 million new cancer cases and 0.6 million cancer deaths estimated in 2018.¹ The annual numbers of new cancer cases and deaths are expected to increase, reaching 2.3 million and 1.0 million, respectively, in 2040.² The direct medical cost associated with cancer care was estimated to increase from \$124 billion in 2010 to

\$173 billion in 2020, a 17% increase over 10 years.³ Reducing cancer burdens through effective prevention strategies has long been an overarching goal for public health policies in the United States.

Suboptimal diet is well known to be associated with the risk of cancer. Strong evidence from systematic reviews suggests that a high consumption of processed and red meats and a low consumption of whole grains and

dairy products are associated with an increased risk of colorectal cancer, and a low consumption of fruits and vegetables is associated with an increased risk of cancer in the oral cavity, pharynx, and larynx.^{4,5} Importantly, obesity has been recognized as a risk factor for 13 types of cancers.⁶ Sugar-sweetened beverage (SSB) consumption can increase the risk of obesity-associated cancers by contributing to weight gain and obesity.^{7,8} We have previously

estimated that more than 80 000 new cancer cases among US adults each year are attributable to suboptimal intake of these dietary factors.⁹ Diet-associated cancers are likely to contribute to a substantial economic burden given the high costs of cancer care. Optimizing dietary intake could be a cost-effective strategy for cancer prevention, yet the economic burden of diet-attributable cancers has not been quantified.

In addition, cancer disproportionately affects individuals of low socioeconomic status and non-Hispanic Blacks in the United States, who bear a higher rate of cancer incidence and death for many cancers than other population subgroups.¹⁰ Meanwhile, dietary disparities have persisted or worsened for most dietary components among US adults despite an overall modest improvement in Americans' diet in the past 10 to 15 years.^{11–13} For example, when low-income adults who participated in the Supplemental Nutrition Assistance Program (SNAP) were compared with higher-income individuals, SNAP participants experienced no improvements in diet quality from 2003 to 2013, whereas diet quality significantly improved among higher-income individuals.¹² Non-Hispanic Blacks had a worse diet quality and smaller improvement in diet quality over time compared with non-Hispanic Whites.^{11,13} Interestingly, the racial/ethnic disparities in colorectal cancer incidence and mortality parallel the racial/ethnic disparities in diet quality among US adults.^{10,14,15} Dietary disparities could contribute to substantial cancer disparities in the United States. In the present study, we aimed to quantify disparities in the health and economic burdens of cancer attributable to suboptimal diet among US adults. Such

findings can inform priority areas in public health strategies to improve diet and reduce diet-attributable cancer disparities in the United States.

METHODS

We used a cohort state-transition model, the Dietary and Cancer Outcome Model,¹⁶ to estimate the health and economic burdens of cancer attributable to suboptimal diet among US adults over a lifetime in a closed cohort (Figure A and Method A, available as supplements to the online version of this article at <http://www.ajph.org>). Starting from a cancer-free representative population of US adults in 2017, the model simulated the development and progression of cancer over a lifetime by transitions of health states and tracked numbers of cancer cases and cancer deaths, and associated direct medical and indirect societal costs that occurred annually in the cohort. To estimate diet-associated cancer burdens, the model compared the incremental difference in cancer burdens between the current and optimal dietary intake scenarios. Diet-associated cancer burdens were estimated in population subgroups stratified by age, gender, race/ethnicity, education, income, and SNAP participation, and combined for estimates among US adults. Health outcomes and economic costs were discounted at 3% annually as recommended by the Second Panel on Cost-Effectiveness in Health and Medicine.¹⁷

Study Population

Demographics were obtained on non-institutionalized US adults aged 20 years or older based on the 2 most recent cycles of the National Health and Nutrition Examination Survey

(NHANES) (2015–2016 and 2017–2018; Table A, available as a supplement to the online version of this article at <http://www.ajph.org>). Population subgroups were jointly stratified by age (20–44, 45–54, 55–64, and 65 years or older), gender (men and women), race/ethnicity (non-Hispanic Whites, non-Hispanic Blacks, Hispanics, and others), education (< high school, high school, some college, and college graduate or above), income (family income to poverty ratio [FIPR], calculated by using the poverty guideline by the Department of Health and Human Services, of < 1.3, 1.3–2.9, and ≥ 3), and SNAP participation status (SNAP participants, SNAP-eligible nonparticipants, and SNAP ineligible individuals). Information on race/ethnicity was self-reported according to fixed categories; Asian and other racial/ethnic groups were combined into 1 group because of their small sample sizes.¹⁸

Current and Optimal Dietary Intakes

Seven dietary factors representing the suboptimal diet (a low consumption of whole grains, dairy products, fruits, and vegetables, and a high consumption of red meats, processed meats, and SSBs) were selected on the basis of evidence from systematic reviews performed by the World Cancer Research Fund/American Institute for Cancer Research (WCRF/AICR) and others showing “convincing” or “probable” evidence of association with cancer risk.⁸ We estimated current intakes of these dietary factors by using two 24-hour dietary recalls per person from NHANES cycle 2015–2016 and 2017–2018, which provided the most recent dietary intake data of nationally representative US adults. The NHANES used the US

Department of Agriculture Automated Multiple-Pass Method to enhance complete and accurate recall of all foods and beverages consumed in the previous day and reduce respondent burden across all cycles.¹⁹ We performed energy adjustment to reduce measurement errors associated with self-reported dietary intake estimates.²⁰ The estimated mean consumption incorporated sampling weights to account for the complex sampling design and ensure national representativeness (Table B, available as a supplement to the online version of this article at <http://www.ajph.org>). We characterized the optimal intake of each dietary factor based on the consumption level associated with lowest disease risk in meta-analyses of clinical end points, assessed by the Global Burden of Disease 2010.²¹

Diet–Cancer Associations

To estimate cancer risks attributable to suboptimal diet, we incorporated both direct diet–cancer etiologic effects and the body mass index (BMI; defined as weight in kilograms divided by the square of height in meters)–mediated associations between diet and cancer. We obtained the relative risk estimates for direct diet–cancer etiologic effects from meta-analysis performed by WCRF/AICR,^{4,5} which included prospective cohort studies with limited evidence of bias from confounders, where the associations were multivariable adjusted and independent of BMI (Table C and Method B, available as supplements to the online version of this article at <http://www.ajph.org>). The long-term etiologic effects of dietary factors on BMI were estimated based on multivariable-adjusted pooled analysis from 120 977 US men and women from 3 prospective cohort studies

(Table D, available as a supplement to the online version of this article at <http://www.ajph.org>).^{7,9} We obtained effects of elevated BMI on cancer from meta-analyses of prospective cohort studies conducted by the International Agency for Research on Cancer⁶ and WCRF/AICR (Table E, available as supplement to the online version of this article at <http://www.ajph.org>).⁸

Cancer Incidence and Survival

We obtained cancer incidence rates for the 15 diet-related cancer types (i.e., colorectal, oral cavity or pharynx, larynx, corpus uteri, kidney, breast, liver, stomach, esophagus, pancreas, prostate, thyroid, gallbladder, ovary, and multiple myeloma) in 2017 from the Centers for Disease Control and Prevention's National Program of Cancer Registries and the Surveillance, Epidemiology, and End Results program (SEER; Method C and Table F, available as supplements to the online version of this article at <http://www.ajph.org>).²² To account for underlying trends in cancer incidence for each cancer type, we estimated the average annual percent change in age-adjusted incidence rates from 2008 to 2017, and then applied that to the baseline incidence rates to project future trends in cancer incidence in 2018 and beyond.³ We estimated the annual probability of dying from cancer based on the 5-year relative survival in 2016 obtained from SEER (Method C and Table G, available as supplements to the online version of this article at <http://www.ajph.org>).

Economic Costs

We obtained direct medical costs for cancer care by age (< 65 and ≥ 65 years),

gender, and phase of cancer care (initial, continuing, and end year of life) from National Cancer Institute's Cancer Prevalence and Cost of Care Projections.³ Indirect societal costs including productivity loss costs and patient time costs associated with cancer care were obtained from published estimates based on the Medical Expenditure Panel Survey data.^{23,24} All costs were inflated to 2017 US dollars using the Personal Health Care index (Method D and Table H, available as supplements to the online version of this article at <http://www.ajph.org>).

Uncertainty Analysis

We incorporated uncertainties in each model input parameter using probabilistic sensitivity analyses with 1000 simulations. We derived corresponding 95% uncertainty intervals (UIs) from the 2.5th and 97.5th percentiles of 1000 estimates. We conducted statistical analyses in R, version 3.6.1 (R Foundation, Vienna, Austria).

RESULTS

Among US adults over a lifetime, suboptimal intakes of 7 dietary factors were estimated to cause 3.04 (95% UI = 2.88, 3.20) million new cancer cases, 1.74 (95% UI = 1.65, 1.84) million cancer deaths (Table 1), and \$253.69 (95% UI = \$241.54, \$266.54) billion direct medical costs associated with cancer care (Table 2), accounting for 7.4% of new cancer cases, 7.7% of cancer deaths, and 7.8% of direct medical costs of these 15 cancers in the United States. These diet-attributable cancers also cost \$113.89 (95% UI = \$108.21, \$119.86) billion in productivity losses and \$16.10 (95%

TABLE 1— Estimated Cancer Burden Attributable to Suboptimal Diet Among US Adults Over a Lifetime by Population Subgroups: 2015–2018

	New Cancer Cases		Cancer Deaths	
	No., Median (95% UI)	No. per 100 000 Population, Median (95% UI)	No., Median (95% UI)	No. per 100 000 Population, Median (95% UI)
All US adults	3 040 000 (2 880 000, 3 200 000)	1 290 (1 230, 1 360)	1 740 000 (1 650 000, 1 840 000)	742 (704, 782)
Gender				
Women	1 400 000 (1 310 000, 1 490 000)	1 150 (1 070, 1 220)	767 000 (714 000, 821 000)	629 (585, 673)
Men	1 640 000 (1 510 000, 1 760 000)	1 450 (1 340, 1 560)	977 000 (906 000, 1 050 000)	863 (801, 928)
Age, y				
20–44	1 610 000 (1 490 000, 1 750 000)	1 550 (1 430, 1 680)	992 000 (919 000, 1 080 000)	953 (883, 1 030)
45–54	564 000 (514 000, 619 000)	1 390 (1 270, 1 530)	308 000 (281 000, 338 000)	761 (694, 835)
55–64	503 000 (453 000, 549 000)	1 180 (1 060, 1 290)	267 000 (239 000, 292 000)	624 (560, 682)
≥ 65	357 000 (323 000, 394 000)	747 (675, 824)	178 000 (161 000, 195 000)	371 (335, 408)
Race/ethnicity				
Non-Hispanic Black	377 000 (348 000, 406 000)	1 400 (1 300, 1 510)	245 000 (226 000, 266 000)	915 (842, 992)
Hispanic	485 000 (441 000, 528 000)	1 330 (1 210, 1 450)	307 000 (277 000, 337 000)	840 (757, 923)
Other	265 000 (243 000, 286 000)	1 130 (1 040, 1 220)	151 000 (138 000, 166 000)	645 (587, 707)
Education				
< high school	398 000 (366 000, 430 000)	1 320 (1 210, 1 420)	242 000 (222 000, 263 000)	801 (733, 869)
High school or GED	766 000 (706 000, 829 000)	1 360 (1 250, 1 470)	453 000 (415 000, 493 000)	802 (735, 873)
Some college	987 000 (911 000, 1 070 000)	1 330 (1 230, 1 430)	583 000 (535 000, 633 000)	784 (720, 852)
College or above	847 000 (775 000, 917 000)	1 140 (1 050, 1 240)	495 000 (449 000, 537 000)	668 (607, 725)
Income ^a				
FIPR < 1.3	668 000 (614 000, 725 000)	1 380 (1 270, 1 500)	408 000 (374 000, 444 000)	845 (773, 918)
FIPR 1.3–2.9	910 000 (835 000, 982 000)	1 310 (1 200, 1 410)	543 000 (494 000, 591 000)	780 (710, 850)
FIPR ≥ 3.0	1 420 000 (1 310 000, 1 520 000)	1 210 (1 110, 1 300)	824 000 (757 000, 888 000)	703 (646, 757)
SNAP participation status ^b				
SNAP participants	596 000 (546 000, 649 000)	1 460 (1 340, 1 590)	365 000 (333 000, 400 000)	894 (815, 980)
SNAP-eligible nonparticipants	285 000 (261 000, 310 000)	1 310 (1 200, 1 430)	171 000 (156 000, 188 000)	788 (717, 867)
SNAP-ineligible individuals	2 090 000 (1 930 000, 2 250 000)	1 210 (1 120, 1 300)	1 220 000 (1 130 000, 1 320 000)	709 (652, 767)

Note. FIPR = family income to poverty ratio; GED = general equivalency diploma; SNAP = Supplemental Nutrition Assistance Program; UI = uncertainty interval. The federal poverty threshold was according to US Department of Health and Human Services.

^aThe ratio of family income to the federal poverty threshold, adjusting for household size. For reference, the 2017 federal poverty threshold for a family of 4 was \$24 600 per year, according to the US Department of Health and Human Services.

^bSNAP participants are those reporting having ever received household SNAP benefits in the past 12 years in the National Health and Nutrition Examination Survey; SNAP-eligible nonparticipants refer to those who are income eligible for SNAP (FIPR < 1.3) while not reporting SNAP participation in the past 12 years; SNAP-ineligible individuals refer to those with FIPR ≥ 1.3.

UI = \$15.30, \$16.92) billion in patient time costs. Among all diet-attributable cancer burdens, 72.0% of new cancer cases (2.09 million; 95% UI = 1.95, 2.22), 71.3% of cancer deaths (1.24 million; 95% UI = 1.16, 1.31; Table I, available as a supplement to the

online version of this article at <http://www.ajph.org>), and 72.8% of direct medical costs (\$184.80 billion; 95% UI = \$173.58, \$194.92 billion; Table J, available as a supplement to the online version of this article at <http://www.ajph.org>) were attributable to

direct diet–cancer etiologic effects, and the remainder were attributable to dietary effects mediated through obesity (Table K and Table L, available as supplements to the online version of this article at <http://www.ajph.org>).

TABLE 2— Estimated Economic Costs of Cancer Attributable to Suboptimal Diet Among US Adults Over a Lifetime by Population Subgroups: 2015–2018

	Direct Medical Costs		Productivity Loss Costs		Patient Time Costs	
	Total Costs, Billion \$, Median (95% UI)	Costs per 100 000 Population, Million \$, Median (95% UI)	Total Costs, Billion \$, Median (95% UI)	Costs per 100 000 Population, Million \$, Median (95% UI)	Total Costs, Billion \$, Median (95% UI)	Costs per 100 000 Population, Million \$, Median (95% UI)
All US adults	253.69 (241.54, 266.54)	107.88 (102.71, 113.34)	113.89 (108.21, 119.86)	48.43 (46.01, 50.97)	16.1 (15.30, 16.92)	6.84 (6.50, 7.19)
Gender						
Women	108.20 (101.23, 114.82)	88.65 (82.95, 94.08)	54.54 (51.00, 58.19)	44.69 (41.79, 47.68)	7.93 (7.46, 8.46)	6.50 (6.11, 6.93)
Men	145.34 (135.40, 156.31)	128.48 (119.07, 138.18)	59.28 (55.02, 63.91)	52.40 (48.64, 56.50)	8.17 (7.53, 8.8)	7.22 (6.66, 7.78)
Age, y						
20–44	121.22 (111.53, 131.50)	116.42 (107.11, 126.29)	55.97 (51.26, 60.80)	53.75 (49.23, 58.39)	6.83 (6.25, 7.43)	6.56 (6.01, 7.14)
45–54	51.92 (47.11, 56.91)	128.31 (116.43, 140.67)	24.27 (22.03, 26.64)	59.99 (54.45, 65.85)	3.53 (3.19, 3.88)	8.71 (7.89, 9.59)
55–64	47.20 (42.53, 51.87)	110.46 (99.53, 121.38)	20.84 (18.78, 22.91)	48.77 (43.95, 53.61)	3.53 (3.17, 3.88)	8.26 (7.41, 9.09)
≥ 65	33.18 (29.74, 36.64)	69.34 (62.16, 76.58)	12.91 (11.60, 14.28)	26.97 (24.25, 29.85)	2.23 (2.00, 2.47)	4.65 (4.18, 5.16)
Race/ethnicity						
Non-Hispanic White	161.09 (149.36, 173.20)	108.57 (100.67, 116.74)	73.92 (68.45, 79.6)	49.82 (46.13, 53.65)	10.64 (9.87, 11.45)	7.17 (6.66, 7.72)
Non-Hispanic Black	32.01 (29.52, 34.39)	119.33 (110.05, 128.22)	13.08 (12.16, 14.08)	48.75 (45.34, 52.49)	1.78 (1.65, 1.90)	6.62 (6.16, 7.07)
Hispanic	38.40 (35.08, 41.64)	105.18 (96.10, 114.07)	16.67 (15.18, 18.14)	45.66 (41.58, 49.68)	2.30 (2.11, 2.48)	6.29 (5.78, 6.79)
Other	22.09 (20.35, 23.82)	94.13 (86.73, 101.51)	10.25 (9.37, 11.02)	43.70 (39.94, 46.98)	1.38 (1.27, 1.48)	5.88 (5.43, 6.29)
Education						
< high school	33.54 (30.99, 36.08)	110.73 (102.32, 119.11)	14.20 (13.15, 15.27)	46.89 (43.43, 50.41)	2.02 (1.88, 2.17)	6.66 (6.19, 7.16)
High school or GED	64.23 (59.45, 69.28)	113.77 (105.30, 122.7)	28.02 (25.90, 30.12)	49.62 (45.87, 53.36)	4.00 (3.71, 4.29)	7.08 (6.57, 7.61)
Some college	82.07 (75.84, 87.92)	110.39 (102.02, 118.26)	36.43 (33.78, 39.15)	49.00 (45.44, 52.67)	5.12 (4.76, 5.49)	6.89 (6.41, 7.39)
College or above	71.46 (65.50, 77.37)	96.48 (88.43, 104.45)	32.00 (29.38, 34.71)	43.20 (39.66, 46.86)	4.47 (4.14, 4.83)	6.04 (5.59, 6.52)
Income ^a						
FIPR < 1.3	54.45 (50.23, 58.60)	112.68 (103.93, 121.25)	23.93 (22.08, 26.01)	49.51 (45.69, 53.81)	3.32 (3.08, 3.58)	6.87 (6.36, 7.42)
FIPR 1.3–2.9	75.22 (68.95, 81.38)	108.08 (99.07, 116.94)	32.91 (30.24, 35.65)	47.30 (43.45, 51.23)	4.59 (4.25, 4.94)	6.60 (6.11, 7.10)
FIPR ≥ 3.0	121.40 (111.64, 129.94)	103.54 (95.22, 110.83)	53.86 (49.54, 57.96)	45.94 (42.25, 49.44)	7.65 (7.07, 8.21)	6.52 (6.03, 7.00)
SNAP participation ^b						
SNAP participants	47.86 (43.99, 51.87)	117.26 (107.76, 127.08)	21.14 (19.45, 22.97)	51.78 (47.65, 56.27)	2.91 (2.68, 3.16)	7.12 (6.57, 7.73)
SNAP-eligible nonparticipants	23.39 (21.43, 25.43)	107.55 (98.55, 116.91)	10.27 (9.38, 11.15)	47.21 (43.14, 51.28)	1.42 (1.31, 1.54)	6.55 (6.02, 7.09)
SNAP-ineligible individuals	177.26 (164.24, 189.96)	102.7 (95.16, 110.06)	78.33 (72.51, 84.12)	45.38 (42.01, 48.74)	11.06 (10.26, 11.84)	6.41 (5.95, 6.86)

Note. FIPR = family income to poverty ratio; GED = general equivalency diploma; SNAP = Supplemental Nutrition Assistance Program; UI = uncertainty interval. The federal poverty threshold was according to US Department of Health and Human Services.

^aThe ratio of family income to the federal poverty threshold, adjusting for household size. For reference, the 2017 federal poverty threshold for a family of 4 was \$24 600 per year, according to the US Department of Health and Human Services.

^bSNAP participants are those reporting having ever received household SNAP benefits in the past 12 years in the National Health and Nutrition Examination Survey; SNAP-eligible nonparticipants refer to those who are income eligible for SNAP (FIPR < 1.3) while not reporting SNAP participation in the past 12 years; SNAP-ineligible individuals refer to those with FIPR ≥ 1.3.

Disparities by Age and Gender

Per 100 000 population, young adults (aged 20–44 years) had a higher number of diet-attributable cancer cases (difference = 803; 95% UI = 656, 952) and cancer deaths (difference = 582; 95% UI = 500, 668) than older adults (≥ 65 years; Table 1). Similar age patterns were observed for all cancer types (Figure 1 and Table M, available as a supplement to the online version of this article at <http://www.ajph.org>) and all dietary factors (Figure 2 and Table N, available as a supplement to the online version of this article at <http://www.ajph.org>). Men had a higher number of diet-attributable cancer cases (difference = 300; 95% UI = 169, 434) and cancer deaths (difference = 234; 95% UI = 157, 315) per 100 000 population than women, overall and across cancer types except for female cancers (e.g., female breast, endometrial, and ovary cancers). Men also had a higher diet-attributable cancer burden than women for nearly all dietary factors except for high SSB consumption, which contributed to a higher number of cancer cases in women than in men (11; 95% UI = –10, 43).

Disparities by Race/Ethnicity

Non-Hispanic Blacks had more diet-attributable cancer cases (difference = 110; 95% UI = –21, 265) and cancer deaths (214; 95% UI = 126, 309) per 100 000 population than non-Hispanic Whites. Non-Hispanic Blacks also had a higher number of diet-attributable cancer burden for most cancer types, with the largest differences seen for colorectal cancers. However, non-Hispanic Whites had a higher

number of diet-attributable cases and deaths for oral cavity, pharynx, or larynx cancers than non-Hispanic Blacks. By dietary factors, non-Hispanic Blacks had a higher number of diet-attributable cancer burden than non-Hispanic Whites for most of the dietary factors, with greater differences attributable to low consumption of dairy and whole grains and high consumption of processed meats and SSBs.

Disparities by Education, Income, and SNAP Status

Compared with those with a college graduate or above level of education, individuals with a lower than college graduate level of education had a higher number of diet-attributable cancer cases (e.g., difference for less than high school vs college graduate = 180; 95% UI = 33, 301) and cancer deaths (132; 95% UI = 46, 219) per 100 000 population. Similar disparities by education were observed for nearly all cancer types, with largest differences seen for colorectal cancer and for nearly all dietary factors with the largest differences attributable to high consumption of SSBs.

Individuals with a low level of family income (FIPR < 1.3) had a higher number of diet-attributable cancer cases (difference = 170; 95% UI = 24, 322) and cancer deaths (difference = 142; 95% UI = 49, 238) per 100 000 population than higher-income individuals (FIPR ≥ 3). Similar disparities were observed across cancer types and dietary factors, with only a few exceptions. Similarly, we observed a higher number of diet-attributable cancer cases and cancer deaths among SNAP participants than eligible nonparticipants (difference = 100; 95% UI = 15, 330) and SNAP-ineligible individuals (difference = 250, 95% UI = 106, 402) for nearly all cancer types

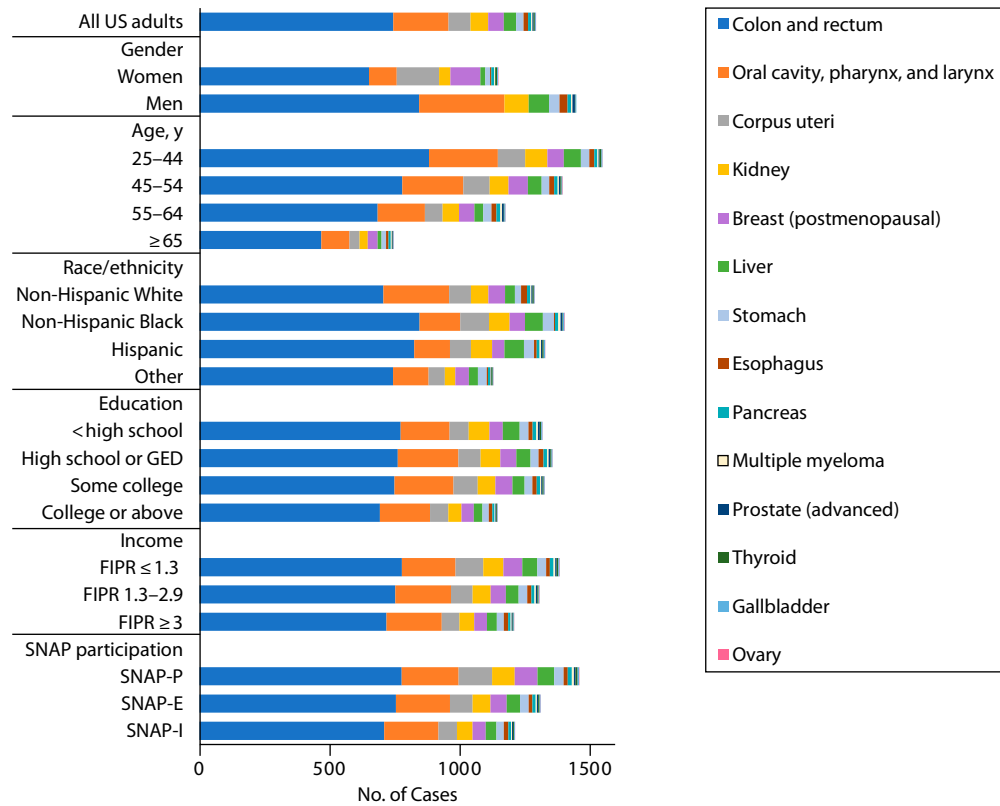
and each dietary factor. The largest disparities by family income and SNAP participation status were attributable to high consumption of SSBs and low consumption of whole grains.

Obesity vs Direct Diet-Cancer Associations

The diet-attributable cancer cases and deaths mediated through obesity were higher among young versus older adults per 100 000 population (difference of cases = 292; 95% UI = 236, 350), women versus men (difference of cases = 128; 95% UI = 78, 182), non-Hispanic Blacks versus non-Hispanic Whites (difference of cases = 95; 95% UI = 21, 160), individuals with low versus higher levels of education (difference of cases = 91; 95% UI = 29, 154) or income (difference of cases = 140; 95% UI = 69, 207), and SNAP participants versus SNAP-ineligible individuals (difference of cases = 162; 95% UI = 87, 242; Table I and Table J). We observed similar disparities for cancer burdens attributable to direct diet-cancer associations, except that men had a higher direct diet-attributable cancer burden than women (Table K and Table L).

Disparities in Economic Costs

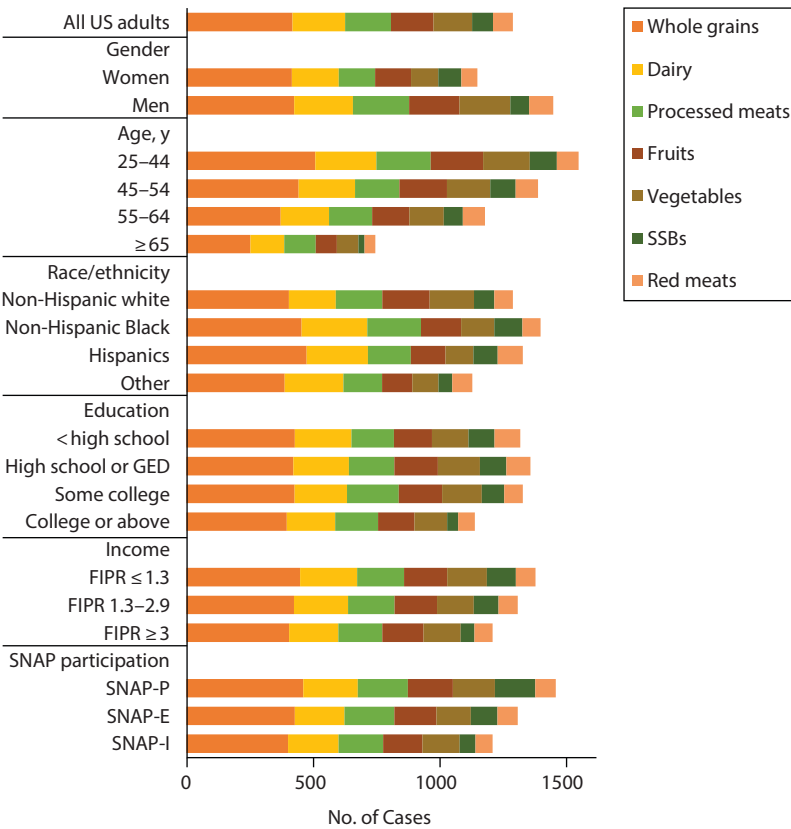
We observed similar disparities for diet-attributable cancer costs for direct medical costs, productivity loss costs, and patient time costs. For example, the direct medical costs of diet-attributable cancers (million per 100 000 population) were higher in young (aged 20–44 years) versus older adults (aged ≥ 65 years; difference = 46.91; 95% UI = 35.69, 59.03), men versus women (difference = 39.90;



	Colon and rectum	Oral Cavity, Pharynx, and Larynx	Corpus Uteri	Kidney	Breast (Postmenopausal)	Liver	Stomach	Esophagus	Pancreas	Multiple Myeloma	Prostate (Advanced)	Thyroid	Gallbladder	Ovary
All US adults	744	212	84	68	60	47	29	17	12	5	5	5	3	1
Gender														
Women	651	107	162	44	115	19	19	4	11	4	0	7	4	2
Men	843	327	0	94	0	78	40	30	13	6	10	2	3	0
Age, y														
25-44	881	264	105	85	63	65	33	18	12	5	4	7	4	1
45-54	779	234	101	72	74	53	29	19	13	5	6	4	4	1
55-64	684	181	68	63	59	34	32	19	15	6	7	3	4	1
≥65	468	108	39	31	38	15	17	9	9	4	4	1	2	1
Race/ethnicity														
Non-Hispanic White	706	253	83	67	64	38	23	23	13	4	5	5	3	1
Non-Hispanic Black	845	156	110	79	59	69	43	4	12	9	6	2	7	1
Hispanic	824	137	81	81	48	75	39	10	11	5	5	6	5	1
Other	743	136	62	40	53	35	33	5	7	3	2	5	3	1
Education														
<high school	772	188	74	79	52	64	34	16	14	7	7	5	4	1
High school or GED	761	233	85	76	62	54	31	18	14	6	6	5	4	1
Some college	749	226	93	68	65	46	30	17	12	5	5	5	4	1
College or above	692	193	70	51	48	33	24	14	8	3	3	3	2	1
Income														
FIPR ≤ 1.3	777	206	106	78	72	57	34	15	14	7	6	6	5	1
FIPR 1.3-2.9	752	214	83	69	58	49	33	16	12	5	5	5	4	1
FIPR ≥ 3	718	212	68	57	48	38	27	17	10	4	4	3	3	1
SNAP participation														
SNAP-P	776	218	128	87	87	64	36	16	16	8	6	8	5	2
SNAP-E	754	207	87	69	62	51	33	15	12	5	5	5	4	1
SNAP-I	709	208	71	60	50	41	29	17	10	4	4	4	3	1

FIGURE 1— Number of Diet-Attributable Cancer Cases by Cancer Types Among Key Population Subgroups of US Adults Over a Lifetime: 2015–2018

Note. FIPR = family income to poverty ratio; GED = general equivalency diploma; SNAP = Supplemental Nutrition Assistance Program. The federal poverty threshold was according to US Department of Health and Human Services.



	Whole Grains	Dairy	Processed Meats	Fruits	Vegetables	SSBs	Red Meats
All US adults	420	207	181	168	153	83	78
Gender							
Women	417	186	143	142	108	91	63
Men	426	232	222	198	201	74	95
Age, y							
25-44	510	241	215	207	185	106	86
45-54	443	223	175	188	171	100	89
55-64	373	191	170	147	136	75	87
≥65	253	134	125	83	85	24	43
Race/ethnicity							
Non-Hispanic white	405	185	183	187	176	79	73
Non-Hispanic Black	456	259	212	158	131	111	72
Hispanic	475	242	169	137	111	95	100
Other	389	232	153	119	104	54	79
Education							
<high school	429	224	168	150	145	103	102
High school or GED	423	220	180	172	166	104	96
Some college	428	207	204	171	157	88	74
College or above	397	191	170	143	130	43	66
Income							
FIPR ≤1.3	450	225	185	170	157	115	78
FIPR 1.3-2.9	425	214	184	168	145	98	76
FIPR ≥3	406	195	173	163	146	55	71
SNAP participation							
SNAP-P	462	215	199	178	166	159	81
SNAP-E	429	196	197	166	136	106	81
SNAP-I	402	199	176	156	147	62	67

FIGURE 2— Number of Diet-Attributable Cancer Cases by Dietary Factors Among Key Population Subgroups of US Adults Over a Lifetime: 2015–2018

Note. FIPR = family income to poverty ratio; GED = general equivalency diploma; SNAP = Supplemental Nutrition Assistance Program; SSB = sugar-sweetened beverage. The federal poverty threshold was according to US Department of Health and Human Services.

95% UI = 29.47, 50.49), non-Hispanic Blacks versus non-Hispanic Whites (difference = 10.69; 95% UI = −.55, 22.84), individuals with low versus higher levels of education (difference = 14.50; 95% UI = 3.53, 25.29) or income (difference = 9.14; 95% UI = 2.39, 20.68), and SNAP participants versus ineligible individuals (difference = 14.75; 95% UI = 2.20, 25.78). However, middle-aged adults (aged 45–54 years) had higher diet-attributable costs than either younger (< 45 years) or older (≥ 55 years) adults. These disparity patterns were similarly observed across cancer types and dietary factors (Table O and Table P, available as supplements to the online version of this article at <http://www.ajph.org>).

DISCUSSION

Based on a nationally representative simulation model, suboptimal diet was estimated to contribute to 3.04 million new cancer cases, 1.74 million cancer deaths, and \$254 billion direct medical costs associated with cancer care among US adults over a lifetime. The health and economic burdens of diet-attributable cancers were higher among men, younger adults, racial/ethnic minorities, individuals with lower education and income attainments, and SNAP participants, compared with their counterpart groups. The largest disparities in diet-attributable cancer were associated with high consumption of SSBs and low consumption of whole grains. To our best knowledge, this study is among the first to quantify disparities in both the health and economic burdens of cancer attributable to suboptimal diet in the United States.

Public Health Implications

The higher diet-attributable cancer burdens among non-Hispanic Blacks, individuals with lower education and income attainments, and SNAP participants reflects both a higher cancer risk and a worse diet quality among these population subgroups.^{10–12} These findings suggest that targeted nutrition interventions among these population subgroups can potentially reduce diet-attributable cancer disparities. In addition, our study revealed that the highest diet–cancer disparities were attributable to high consumption of SSBs and processed meats and low consumption of whole grains and dairy. These could be priority dietary targets for behavior change and policy strategies to reduce cancer disparities in the United States. Policy options may include expanding SNAP to include financial incentives for purchasing whole grains, fruits, and vegetables and disincentives to discourage the consumption of SSBs and processed meats.^{25,26} Other relevant policy actions may include improving the availability, affordability, and accessibility of healthy foods in low-income and racial/ethnic minority communities, schools, and workplaces.²⁷

The gender difference in diet-attributable cancer burdens is also worth attention. The overall higher diet-attributable cancer burdens in men than women may reflect a worse diet quality in men than women.^{11,13} For example, the Healthy Eating Index 2015, a diet quality index that measures adherence to the 2015–2020 Dietary Guidelines for Americans, was 5% (3 percentage points) lower in men than women.¹³ However, women had a higher diet-attributable cancer burden mediated through obesity than men. It

is possible that excessive body weight has a larger impact on female cancers. In accordance with this finding, our results also revealed that high SSB consumption contributed to more cancer cases in women than in men. The higher diet-related cancer burdens estimated among younger than older adults were primarily attributable to the longer length of follow-up of younger adults in a closed cohort.

Consumption of SSBs contributed to the largest diet-attributable cancer disparities in the United States, by age, race, and socioeconomic status. Although SSB consumption had declined by 28% since 1999 among US adults, level of consumption remains high, especially among young adults, racial/ethnic minorities, and socioeconomically disadvantaged groups.¹² Reducing SSB consumption through relevant policy actions, including taxes, warning labels, or Nutrition Facts Panel labeling of added sugars^{28,29} could be effective strategies for reducing diet-attributable cancer disparities. Although restricting SSB purchases for SNAP participants has been debated,³⁰ a combination of financial disincentives for SSBs and other less healthful foods plus incentives for a range of healthful foods may help reduce disparities while still preserving choice.³¹

Low whole grain consumption also contributed to substantial diet–cancer disparities in the United States. Despite modest recent increase in whole grain consumption, current levels of 1 serving a day remain far below the recommended intake of 3 servings per day.¹¹ In 2015 to 2016, US adults consumed only 2.7% of calories from whole grains, compared with 15.9% of calories from refined grains.¹³ Potential barriers for increasing whole grain consumption include the lack of public awareness on

health benefits of whole grains, lack of knowledge to identify whole grain products, and absence of standardized definitions and labeling to increase awareness and healthier choices.³² Efforts are needed to address these challenges,^{9,32} including incorporating additional financial incentives for healthy whole grain products (rather than only fruits and vegetables) in SNAP.³¹

By cancer type, optimizing dietary intake could be particularly important for reducing disparities related to colorectal cancer. Colorectal cancer is the third most diagnosed cancer among US men and women¹ and is associated with the largest number of diet-attributable cases.⁹ Colorectal cancer disproportionately affects men, non-Hispanic Blacks, and socioeconomically disadvantaged groups,^{14,15} which runs parallel with the higher colorectal cancer burden attributable to diet observed in this study.

Strengths and Limitations

Our study had several strengths. We used a probabilistic cohort state-transition model, which simulates the transitions among various cancer-related health states along with aging, allowing us to project the health and economic burdens of cancer attributable to diet over a lifetime. Our model incorporated national representative estimates for recent dietary intakes of US adults, national data for cancer incidence and survival, and multivariate-adjusted etiology effects for diet–cancer associations from meta-analysis of cohort studies. In addition to estimating direct diet–cancer associations independent of obesity, our model further incorporated obesity-mediated cancer risks, allowing us to estimate diet-attributable

cancer burdens attributable to obesity-mediated pathways. Our model also accounted for the uncertainty of dietary intake and cancer incidence, allowing estimation of the lower and upper bounds of the plausible effects.

Potential limitations should also be considered. First, self-reported dietary assessment is subject to measurement error. Yet the NHANES is the only nationally representative survey of Americans' diet; dietary intake data collected in NHANES are often used to evaluate dietary intake patterns of US adults and children.¹¹ In addition, the use of multiple-pass method¹⁹ and energy adjustments reduce measurement error.²⁰

Second, we did not incorporate secular trends in diet over time or changes in diet across the life course. If disparities in dietary intake patterns persist or worsen in future years or as people age, the estimated disparities in diet-associated cancer burdens are likely to sustain to later years or become more profound. In addition, the current estimates did not consider the impact of early life diet on cancer outcomes and could have underestimated cancer burdens attributable to suboptimal diet.

Third, the projected rates of cancer incidence and mortality in the current model may not fully capture the influence of Affordable Care Act and other changing policies and factors on cancer outcomes. Disparities in diet-associated health and economic burdens of cancer may decline in future years if the insurance expansion under the Affordable Care Act is likely to reduce cancer disparities in the United States.³³

Fourth, cancer could lead to substantial psychological burdens for the patients and their families. Our model did not include the psychological

burden of cancer because of the lack of well-accepted methods to quantify it;³⁴ the diet-associated cancer burdens could be underestimated.

Fifth, we assumed a 5-year induction period between changes in current dietary intakes to optimal intakes and cancer risks, based on limited empirical evidence of the induction time between diet and various cancers.³⁵ Longer or shorter induction periods could partly alter our findings.

Conclusions

Suboptimal diet contributes to a substantial cancer burden in racial/ethnic minorities, socioeconomically disadvantaged groups, SNAP participants, men, and young adults. High consumption of SSBs and low consumption of whole grains are 2 leading dietary targets for prevention policies to reduce diet-attributable cancer disparities in the United States. These findings underscore the importance of developing and implementing targeted food and nutrition strategies among key population subgroups to reduce cancer disparities in the United States. *AJPH*

ABOUT THE AUTHORS

Lu Wang, Mengxi Du, Frederick Cudhea, Christina Griecci, Dariush Mozaffarian, and Fang Fang Zhang are with the Friedman School of Nutrition Science and Policy, Tufts University, Boston, MA. Dominique S. Michaud is with the Department of Public Health and Community Medicine, School of Medicine, Tufts University.

CORRESPONDENCE

Correspondence should be sent to Fang Fang Zhang, Friedman School of Nutrition Science and Policy, Tufts University, 150 Harrison Ave, Boston, MA 02111 (e-mail: fang_fang.zhang@tufts.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

L. Wang and F. F. Zhang conceptualized the study. L. Wang performed formal data analysis and drafted the article. M. Du and C. Griecci provided key input data used in the simulation model. F. Cudhea contributed to the development of the simulation model. D. Mozaffarian and F. F. Zhang supervised the study. F. F. Zhang acquired funding for conducting the study. M. Du, F. Cudhea, D. S. Michaud, and D. Mozaffarian and F. F. Zhang critically reviewed the article and provided important intellectual content. All authors read and approved the final article.

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CONFLICTS OF INTEREST

The authors declare no competing interests.

HUMAN PARTICIPANT PROTECTION

The study is exempt for ethical review and waived for consent because it used publicly available data with no personal identifiable information.

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Severity of Inpatient Hospitalizations Among Undocumented Immigrants and Medi-Cal Patients in a Los Angeles, California, Hospital: 2019

Annie Ro, PhD, MPH, Helen W. Yang, MD, Senxi Du, BA, Courtney L. Hanlon, MD, MS, and Andrew Shane Young, DO

 See also Ortega et al., p. 1910.

Objectives. To compare the severity of inpatient hospitalizations between undocumented immigrants and Medi-Cal patients in a large safety-net hospital in Los Angeles, California.

Methods. We conducted a retrospective analysis of all 2019 inpatient stays at a Los Angeles hospital (n = 22 480), including patients of all races/ethnicities. We examined 3 measures by using insurance status to approximate immigration status: illness severity, length of hospital stay, and repeat hospitalizations. We calculated group differences between undocumented and Medi-Cal patients by using inverse probability weighted regression adjustment separately for patients aged 18 to 64 years and those aged 65 years and older.

Results. Younger undocumented patients had less severe illness and shorter lengths of stay than their Medi-Cal counterparts. Older undocumented immigrants also had less severe illness, but had similar lengths of stay and were more likely to have repeated hospitalizations.

Conclusions. While existing work suggests that undocumented immigrants could have more severe health care needs on account of their poorer access to medical care, we did not see clear health disadvantages among hospitalized undocumented immigrants, especially younger patients. There were fewer differences between undocumented and Medi-Cal patients who were older. (*Am J Public Health.* 2021;111(11):2019–2026. <https://doi.org/10.2105/AJPH.2021.306485>)

There is growing interest in how immigration status affects the health care access and utilization for 10.5 million undocumented immigrants¹ in the United States. Undocumented immigrants have significantly lower levels of insurance coverage and lower use of the emergency department, and are less likely to have a usual source of care compared with both their documented counterparts and the US-born.^{2,3} These disparities are attributable to lower household incomes and ineligibility for Medicaid

and Medicare, as well as lack of private insurance from employers. Even when health care is available, undocumented immigrants may be reluctant to access it because of the fear of deportation for themselves or their families.^{4,5} Undocumented immigrants also face challenges accessing the limited care that is available to them because of economic barriers and unfamiliarity with the health care system.⁶

Despite worse access to and lower utilization of health care, undocumented immigrants do not seem to

have widespread negative physical health. In a review of 45 studies, Hamilton et al.⁷ found that the large majority did not observe undocumented immigrants to have significantly worse health than their documented counterparts. In some studies, undocumented immigrants seemed to have better health outcomes for conditions such as hypertension, asthma, and other self-reported chronic conditions. A nationally representative sample confirmed the diverging trends between health insurance coverage and health status:

Ro and Van Hook⁸ found that undocumented immigrants had lower odds of being currently insured but also had lower odds of disability and poor or fair self-rated health compared with those born in the United States.

While undocumented immigrants' limited health care access and robust physical health patterns may appear contradictory, these findings rely on self-reported health status in surveys and may capture their health care utilization when individuals are relatively healthy. Undocumented immigrants tend to be younger than their documented counterparts,⁸ suggesting that some chronic health conditions may not have yet emerged or may have not reached the point of requiring serious medical intervention. Undocumented immigrants are also more likely to be positively selected on health, given the costs and risks inherent in unauthorized migration.⁹ Undocumented immigrants' younger age and positive health selection may buoy their general health status, and, as a result, they may not seek regular medical care.

What is unknown, however, is the health status of undocumented immigrants when they reach the point of needing intensive medical care and their health needs when they are in close contact with the health care system. As immigrants stay longer in the United States, their positive health advantage erodes.¹⁰ In the case of undocumented immigrants, initial health advantages can obscure longer-term care needs as their health deteriorates over time. In the general population, those who are uninsured or do not have regular medical care are more likely to enter the health care system in poorer health and have worse outcomes for both overall health and specific diseases.¹¹ Given their overall

lack of medical care access, it is possible that undocumented immigrants will display poorer health outcomes than other groups when they are at the point of needing higher-level care. Moreover, consistent underutilization of preventive services may leave them with more advanced stages of disease when they eventually need medical attention.

The extant literature has provided only limited or dated information on this topic. One study found uninsured undocumented immigrants hospitalized in Florida to have higher case mix indexes (a measure of disease severity) but shorter hospital stays than immigrants with legal status.¹² Among patients in select cities in California and Texas, undocumented immigrants had the same levels of hospitalizations as documented immigrants, with the exception of childbirth.¹³ Treatment of end-stage renal disease among undocumented immigrants has been widely studied as an example of a life-threatening condition that requires intensive medical treatment but is highly mediated by access to health care. Undocumented immigrants with end-stage renal disease are more likely to receive emergency-only hemodialysis than standard hemodialysis,¹⁴ which is associated with increased mortality, health care utilization, and costs.^{15,16}

In this study, we explored the severity of hospitalizations among undocumented immigrants of all races/ethnicities who have required inpatient stays at the largest safety net hospital in Los Angeles County, California. Los Angeles County has the largest population of undocumented immigrants in the country; there are nearly 880 000 undocumented immigrants, representing a wide number of countries of origin.¹⁷ We used insurance status to

approximate immigration status, leveraging the county's unique health plans for undocumented immigrants to accurately estimate the likely undocumented population. Addressing this knowledge gap will reveal potential unmet needs of undocumented immigrants in the hospital setting during critical periods of illness as well as possible disparities in inpatient health care.

METHODS

The study was a retrospective analysis of all inpatient stays at Los Angeles County and University of Southern California (LAC+USC) Medical Center from January 1, 2019, to December 31, 2019. We merged 2 data sources: (1) the LAC+USC internal electronic medical records system using Cerner PowerIn-sight and (2) Vizient Health System Data, a hospital billing and administrative claims database. All data were de-identified to conform to Health Insurance Portability and Accountability Act requirements.

We examined the data at both the encounter and patient level. For encounter data, there were 29 765 total inpatient hospitalizations at LAC+USC in 2019. We limited the sample to hospitalizations with patients aged 18 years and older and who were full-scope Medi-Cal patients (California's Medicaid program) or undocumented (coding detailed in the "Variables" section). Our final analytic sample consisted of 22 480 inpatient encounters. In our analyses, we separated the sample by age (18–64 years and ≥ 65 years) because of differences in health status at older ages and public insurance coverage (18–64 years = 18 244; ≥ 65 years = 4236). For the older adults, we limited the sample to those who only had Medi-Cal or a combination of

Medi-Cal and Medicare to better isolate a low-income comparison group. For outcomes that used patient-level data (e.g., unique number of patients who were admitted to LAC+USC in 2019), there were 15 876 patients (18–64 years = 12 910; ≥ 65 years = 2966).

Variables

Outcomes. We examined 3 outcomes indicating the severity of the hospitalization: illness severity, length of hospital stay, and repeat hospitalization.

We measured illness severity by the relative risk of mortality, which we calculated with a proprietary algorithm that predicts the risk for mortality based on patient demographics, clinical characteristics, procedures, and comorbidities for each Medicare Severity–Diagnosis Related Group. For example, the risk prediction for a liver transplant (Medicare Severity–Diagnosis Related Group 5 or 6) was a function of hemodialysis, cachexia, complication of transplanted organ or tissue, ventilator on admission day, and type 2 diabetes. The risk for mortality was averaged over all encounters in the hospital for each diagnosis group in 2019. Encounters that were 75% of the mean were coded “lower risk” relative to the mean. Encounters that were within 75% and 125% of the mean were coded “similar risk,” and encounters more than 125% of the mean were considered “higher risk.” We dichotomized the outcome to hospitalizations lower or similar to the mean versus those with higher risk of mortality.

Length of hospital stay was the total number of days of the inpatient admission, with longer encounters indicating sicker patients. This was an encounter-level variable, meaning that each

hospitalization had its own length of stay. To account for outliers, we conducted additional analyses with encounters longer than 21 days removed.

Repeat hospitalization was a patient-level variable and was dichotomized to patients who had 1 inpatient admission in 2019 versus those who had more than 1 admission. We assumed that those with more than 1 admission were sicker and needed more comprehensive care. Though it is possible that repeated hospitalizations might be attributable to unrelated or isolated events (e.g., injuries), it is more likely that repeated hospitalizations result from unresolved or ongoing illness.

Immigration status. We compared undocumented immigrants to full-scope Medi-Cal patients by using insurance status as a proxy to determine immigration status. We coded a patient as having undocumented status if the primary insurance for the encounter was restricted-scope Medi-Cal, which provides health services to low-income Los Angeles County residents who meet the income threshold for Medi-Cal but do not meet immigration status requirements as either US nationals, citizens, or lawful permanent residents. These services include access to county facilities that provide preventive, emergency, diagnostic, specialty, inpatient, and pharmacy services, as well as a local health care program (My Health LA) that offers primary and preventive health care services through community clinic partners.¹⁸

Given the income requirement for restricted-scope Medi-Cal eligibility, we chose full-scope Medi-Cal (hereafter referred to as Medi-Cal) patients as a comparison group of low-income patients who are either US-born or

foreign-born with authorized status. Because of the citizenship requirements for federal health insurance, nearly all of the Medi-Cal patients are either US-born or documented. The other option for a comparison group is individuals on other forms of insurance (e.g., private insurance), but we do not know either the poverty level or the immigration status of these patients.

Covariates. We included age as a continuous variable, race/ethnicity (Hispanic [reference], non-Hispanic White, non-Hispanic Black, non-Hispanic Asian, non-Hispanic other), language (English [reference], Spanish, other), gender (male [reference], female), and homeless status (housed [reference], non-housed). We included language preference to account for reduced English proficiency, which can have an impact on treatment utilization and quality.¹⁹ As the primary safety net hospital in the county, LAC+USC has a substantial homeless patient population. We included homeless status to account for the substantial barriers to regular medical care that unhoused individuals face, as well as their complex health conditions.²⁰ For encounter-level outcomes, we included an indicator of whether the encounter was the only admission for the patient in 2019 (reference) or if it was a repeat hospitalization. For instance, if an encounter was 1 of 2 or more hospitalizations from a patient in 2019, all encounters from that patient were coded “1.” If an encounter was the only hospitalization from a patient in 2019, it was coded “0.”

Analysis

Because of the inherent compositional differences between undocumented

and Medi-Cal patients, we estimated group differences in our outcomes of interest by using inverse probability weighted regression adjustment (IPWRA). This approach weights observations based on the inverse of their conditional probability of “treatment” exposure (i.e., being undocumented). The “treatment” model is estimated first, and the predicted probability of being “treated” (i.e., undocumented) is used to compute the inverse-probability weights. These weights are then applied to the outcome models (i.e., illness severity, length of stay, repeat admission), creating a pseudo-matched sample, such that undocumented and Medi-Cal patients are balanced in their covariates. In conventional multivariate regression analysis, covariates are included in the model to account for compositional differences but each observation is weighted equally. In IPWRA models, observations that have a higher likelihood of being “treated” (i.e., undocumented) are up-weighted while those that have lower likelihood are down-weighted. IPWRA is also known as the doubly robust method because it models both the outcome and propensity for “treatment” within the same framework, and only one needs to be correctly specified to produce unbiased results.²¹

We ran each weighted outcome model on the undocumented group first and then the comparison group. We calculated average treatment effects (ATEs), a common postestimation approach for treatment effects models such as the IPWRA,²² for each of our outcomes. The ATEs take a counterfactual approach and assume that each subject has a pair of potential outcomes: the outcome if they were undocumented or the outcome if they were on Medi-Cal. We calculated the

predicted means for the outcome for each observation assuming they were undocumented (using their own covariate values) and then again for each observation assuming they were on Medi-Cal. For each observation, the difference between the undocumented predicted mean and Medi-Cal predicted mean is called the effect of “treatment.” We averaged this difference across the entire population to calculate the ATE.

For the “treatment” model (i.e., undocumented vs Medi-Cal), we included age, race, language, gender, and homeless status as predictors of undocumented status. For the outcome models (i.e., illness severity, length of stay, repeat admission), we included the same covariates as the “treatment” model as well as whether the encounter was a repeat hospitalization for the illness severity and length-of-stay outcomes. For illness severity and repeat admissions, we conducted logistic regressions, whereas for length of stay, we conducted a linear regression model. We calculated robust standard errors to correct for potential for heteroskedasticity. We did this separately for patients aged 18 to 64 years versus those aged 65 years or older. We conducted our analyses by using Stata version 16 (StataCorp LP, College Station, TX). We provide the coefficients for the IPWRA models in the Tables C, D, and E (available as supplements to the online version of this article at <http://www.ajph.org>).

RESULTS

Table 1 provides descriptive information for our sample at the encounter level. The same descriptive statistics at the patient level are provided in Table A (available as a supplement to the online

version of this article at <http://www.ajph.org>). Nearly one third of encounters were for undocumented patients (31.6%). The mean age for younger patients (18–64 years) was 45.8 years for undocumented patients and 41.6 years for Medi-Cal patients. Among older patients (≥ 65 years), the mean age was 73.1 years for undocumented and 73.8 years for Medi-Cal. Undocumented patients in both age groups were predominantly Hispanic (18–64 years: 91.8% vs ≥ 65 years: 86.5%). Medi-Cal patients in both age groups were also predominantly Hispanic, but comprised a lower proportion (18–64 years: 62.5% vs ≥ 65 years: 57.2%).

Undocumented patients primarily spoke Spanish (18–64 years: 82.5%; ≥ 65 years: 84.8%), whereas Medi-Cal patients had differences in language preference according to age group. Notably, homelessness was nearly 3-fold greater among Medi-Cal patients compared with undocumented patients across age groups (18–64 years: 18.8% vs 6.6%; ≥ 65 years: 13.3% vs 4.8%).

Compared with Medi-Cal patients, undocumented patients had lower illness severity and average length of stay. Among younger patients, 8.6% of undocumented patients had higher-than-average relative mortality risks compared with 11.2% of Medi-Cal patients. Among older patients, 15.8% of undocumented patients had higher-than-average mortality risk compared with 23.1% of Medi-Cal patients. The average length of stay in the hospital was 4.6 days for undocumented and 6.1 days for Medi-Cal among younger patients and 5.2 days and 6.8 days, respectively, among older patients. Among younger patients, inpatient encounters over the study period were comparable between undocumented

TABLE 1— Descriptive Table of Inpatient Encounters in Los Angeles County and University of Southern California Medical Center: 2019

	Age 18–64 y			Age ≥ 65 y		
	Undocumented (n = 6 074), Mean or No. (%)	Medi-Cal (n = 12 170), Mean or No. (%)	Total (n = 18 244), Mean or No. (%)	Undocumented (n = 1 226), Mean or No. (%)	Medi-Cal (n = 3 010), Mean or No. (%)	Total (n = 4 236), Mean or No. (%)
Mean age, y	45.9	41.4	42.9	73.1	73.8	73.6
Race/ethnicity						
Hispanic	5 576 (91.80)	7 609 (62.52)	13 185 (72.27)	1 061 (86.54)	1 723 (57.24)	2 784 (65.72)
Non-Hispanic White	6 (0.10)	508 (4.17)	514 (2.82)	11 (0.90)	88 (2.92)	99 (2.34)
Non-Hispanic Black	75 (1.23)	1 893 (15.55)	1 968 (10.79)	19 (1.55)	443 (14.72)	462 (10.91)
Non-Hispanic Asian	241 (3.97)	461 (3.79)	702 (3.85)	85 (6.93)	370 (12.29)	455 (10.74)
Non-Hispanic other	176 (2.90)	1 699 (13.96)	1 875 (10.28)	50 (4.08)	386 (12.82)	436 (10.29)
Gender						
Female	3 008 (49.52)	4 964 (40.79)	7 972 (43.70)	704 (57.42)	1 386 (46.05)	2 090 (49.34)
Male	3 066 (50.48)	7 206 (59.21)	10 272 (56.30)	522 (42.58)	1 624 (53.95)	2 146 (50.66)
Language						
English	859 (14.14)	8 229 (67.62)	9 088 (49.81)	93 (7.59)	1 106 (36.74)	1 199 (28.31)
Spanish	5 007 (82.43)	3 630 (29.83)	8 637 (47.34)	1 040 (84.83)	1 534 (50.96)	2 574 (60.76)
Other	208 (3.42)	311 (2.56)	519 (2.84)	93 (7.59)	370 (12.29)	463 (10.93)
Homeless	407 (6.70)	2 289 (18.81)	2 696 (14.78)	59 (4.81)	401 (13.32)	460 (10.86)
Illness severity (risk for mortality)						
Below or similar	5 550 (91.37)	10 804 (88.78)	16 354 (89.64)	1 032 (84.18)	2 314 (76.88)	3 346 (78.99)
Above	524 (8.63)	1 366 (11.22)	1 890 (10.36)	194 (15.82)	696 (23.12)	890 (21.01)
Mean length of stay, days	4.5	6.1	5.6	5.2	6.8	6.3
2019 repeat encounter						
First encounter	4 286 (70.56)	8 624 (70.86)	12 910 (70.76)	800 (65.25)	2 166 (72.00)	2 966 (70.00)
Second or higher	1 788 (29.44)	3 546 (29.13)	5 334 (29.24)	426 (34.75)	844 (28.00)	1 270 (30.00)

and Medi-Cal patients, in which approximately 70.7% were first encounters and 29.1% were repeat hospitalizations. However, among older patients, undocumented patients had more repeat hospitalizations compared with Medi-Cal patients (34.8% vs 24.8%).

Group Differences in Inpatient Measures

The results of the IPWRA are presented in Table 2. Among patients aged 18 to 64 years, encounters with undocumented patients had lower illness severity and shorter lengths of stay

than encounters with Medi-Cal patients. The probability of having an encounter with a higher-than-average risk for mortality was 8.4% for undocumented patients and 11.6% for Medi-Cal patients. Encounters for undocumented patients therefore had 3.2% lower probability of a higher-than-average mortality risk compared with encounters for Medi-Cal patients. The average length of stay for undocumented patients' encounters was 1.1 days shorter than for Medi-Cal patients (5 days vs 6.1 days). Among younger patients, the probability of having a repeat hospitalization in 2019 did not

significantly differ between undocumented and Medi-Cal patients. Both groups had a probability of repeat hospitalization that was around 22%.

Among patients aged 65 years and older, encounters for undocumented patients had similarly lower illness severity. Undocumented patients' encounters therefore had 6.5% lower probability for a higher-than-average risk encounter (15.7% vs 22.2%). There was no significant difference in predicted length of stay. However, the probability of having a repeated hospitalization was higher for undocumented patients compared with

TABLE 2— Group Differences Between Undocumented and Medi-Cal Patients in Inpatient Measures by Age Group, Los Angeles County and University of Southern California Inpatient Admissions: 2019

	Illness Severity (Above-Average Risk for Mortality)		Length of Stay, Days		Repeat Patient	
	Predicted Probability (95% CI)	P	Predicted (95% CI)	P	Predicted Probability (95% CI)	P
Aged 18–64 y		< .001		< .05		.97
Undocumented	8.4 (7.3, 9.5)		4.9 (4.2, 5.6)		22.4 (20.0, 24.9)	
Medi-Cal	11.6 (11.0, 12.2)		6.1 (5.9, 6.3)		22.4 (21.4, 23.4)	
Difference	3.2 (1.9, 4.4)		1.2 (0.4, 1.9)		–0.1 (–2.7, 2.6)	
Aged ≥ 65 y		< .001		.33		< .05
Undocumented	15.7 (13.4, 18.1)		6.0 (5.0, 7.0)		26.5 (22.4, 30.6)	
Medi-Cal	22.2 (20.7, 23.7)		6.6 (6.2, 7)		22.0 (20.2, 23.8)	
Difference	6.5 (3.7, 9.3)		0.5 (–0.6, 1.6)		–4.5 (–9.0, –0.1)	

Note. CI = confidence interval. Predicted probabilities calculated from inverse probability weighted regression adjustment models. Covariates for “treatment” model (i.e., undocumented vs Medi-Cal) were age, race, language, gender, and homeless status. Covariates for outcome models (i.e., illness severity, length of stay, repeat admission) were age, race, language, gender, homeless status, and length of stay.

Medi-Cal patients. Undocumented patients had a 26.5% probability of having a repeat hospitalization in 2019 while Medi-Cal patients had a 22.0% probability, with a statistically significant difference in probabilities of 4.5%.

Sensitivity Checks

We conducted several sensitivity checks to confirm the robustness of our findings. We also limited the sample to those who identified as Hispanic/Latino, which was 72% of the sample aged 18 to 64 years and older and 62% of sample aged 65 years and older. For encounter-level outcomes (relative risk of mortality and length of stay), we kept only 1 encounter per patient. The results were qualitatively similar for all of these checks (Table B, available as a supplement to the online version of this article at <http://www.ajph.org>).

For the length-of-stay outcome, we removed outliers whose inpatient stays were 21 days or longer. For encounters with patients aged 18 to 64 years, the difference in predicted length of stay

dropped to 0.5 days but remained significantly different from zero. For patients aged 65 years and older, the difference dropped to 0.1 and remained nonsignificant.

DISCUSSION

We compared the severity of inpatient stays between undocumented and full-scope Medi-Cal patients at the largest safety-net hospital in Los Angeles County. Our results provide insight into the health status of undocumented immigrants at the point of needing inpatient medical care, by using measures of illness severity, length of hospital stay, and repeat hospitalizations. In our study, younger undocumented patients who were hospitalized had less severe illness and spent less time in the hospital compared with Medi-Cal patients. While existing work suggests that undocumented immigrants could have higher risk for poorer inpatient outcomes on account of their limited access to medical care,²³ the younger undocumented patients in this study

did not present with more severe health problems upon hospital admission. Our findings add to those of others that have found positive health trends among undocumented immigrants compared with documented immigrants or US-born counterparts.^{7,8}

These trends were consistent among older undocumented immigrants as well, but older undocumented patients were more likely to have repeat hospitalizations than their Medi-Cal counterparts. Lower illness severity among older undocumented patients, however, suggests that their higher likelihood of repeat hospitalizations may represent a higher willingness to access care rather than comparatively poorer health. While qualitative work has suggested that barriers to receiving regular care can exacerbate older undocumented immigrants' existing chronic conditions,²⁴ we did not find this to be the case from our data.

There are several caveats to our interpretations. First, the Medi-Cal patients at LAC+USC may be a unique, low-income sample. The high

proportion of homelessness, for example, indicates that this population has complex health care needs. The longer length of stay among Medi-Cal patients may also be attributable to discharge planning issues, such as a long wait for other health facilities or a lack of a discharge destination. Thus, the relatively positive outcomes among undocumented patients may not be attributable to any inherent health advantages per se, but rather that the comparison group suffers from serious health conditions. Conversely, our sample of undocumented immigrants included those who had access to county and community facilities via local health plans that provide preventive care. Los Angeles County has actively addressed the health care needs of the undocumented population, which suggests that this undocumented patient population may be receiving regular care that contributes to their better health at the point of hospitalization.

Limitations

We do not have direct information on the immigration status of patients. However, we believe restricted-scope Medi-Cal is a valid approximation of undocumented status. Patients who are not insured at the time of admission but are Medi-Cal eligible are coded under a separate payment source that provides qualified individuals immediate access to temporary Medi-Cal while applying for permanent Medi-Cal or other health coverage (e.g., hospital presumptive eligibility). There may also be immigrants who use restricted-scope Medi-Cal who are not undocumented, such as those on student and work visas, and certain permanent legal residents who have not met the 5-year residency requirement for public

insurance. Alternatively, there may be undocumented immigrants who do not qualify for restricted Medi-Cal because of their high incomes. We believe these comprise a very small number in our sample, however.

We also acknowledge that we lack important variables, such as nativity, country of origin, or ethnic subgroups (e.g., Mexican, Salvadoran, Chinese) and did not control for them in our analysis. We did include language use in our IPWRA models, however, which up-weights individuals in the full-scope Medi-Cal comparison who may be foreign-born and prefer Spanish. We also found similar results when we limited the analyses to Hispanics/Latinos, confirming that these trends are consistent for Hispanic/Latino undocumented immigrants. In addition, we did not have information on baseline health status. While this was indirectly incorporated in our illness severity measures, this is a limitation for length of stay and repeat encounters. Finally, our illness severity outcome might be biased by undiagnosed disease. Yet comprehensive evaluation upon admission often leads to diagnosis of various chronic conditions, which are then captured in the illness severity score. Nevertheless, we acknowledge the limitations of unknown chronicity of medical conditions and adequacy of treatment in the outpatient setting. Future research, including qualitative interviews, would enrich research with respect to nativity status, length of time in the United States, and perceptions around health and health care utilization among undocumented immigrants.

Conclusions and Future Directions

We found that undocumented immigrants in Los Angeles County do not

have poorer health outcomes at the point of hospitalization. Contrary to some popular narratives, undocumented immigrants do not appear to overburden the health care system. We note, however, that the county provides public or subsidized access to health care services for its undocumented residents. While we do not have information on primary care utilization or outpatient care, future research could examine whether local policies contribute to positive hospitalization characteristics among undocumented immigrants or whether our findings hold in other locales with weaker safety-net programs. [AJPH](#)

ABOUT THE AUTHORS

Annie Ro is with the Department of Health, Society, and Behavior at the University of California, Irvine. At the time the study was conducted, Helen W. Yang was a chief resident in Internal Medicine at the Keck School of Medicine of University of Southern California (USC), Los Angeles. Senxi Du is a medical student at the Keck School of Medicine of Medicine. Courtney L. Hanlon is with the Los Angeles County and USC Medical Center, Los Angeles. Andrew Shane Young is with the Division of Geriatric, Hospital, Palliative, and General Internal Medicine, Department of Medicine, Keck School of Medicine.

CORRESPONDENCE

Correspondence should be sent to Annie Ro, PhD, MPH, Anteater Instruction and Research Building, Room 2036, 653 E Peltason Rd, Irvine, CA 92697-3957 (e-mail: annie.ro@uci.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

A. Ro conceptualized the study, conducted the analysis, and wrote the article. H. W. Yang and A. S. Young conceptualized the study, interpreted results, and assisted with article writing. S. Du and C. L. Hanlon interpreted results and assisted

with article writing. All authors reviewed the final draft of the article.

CONFLICTS OF INTEREST

The authors have no conflicts of interests to disclose.

HUMAN PARTICIPANT PROTECTION

All project activities were reviewed and approved by the USC institutional review board (HS-19-00890), which served as a reliance for the University of California Irvine institutional review board. The project did not require informed consent.

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Vaccine Coverage Across the Life Course in Michigan During the COVID-19 Pandemic: January–September 2020

Angela K. Shen, ScD, MPH, Cristi A. Bramer, MPH, Lynsey M. Kimmins, MPH, Robert Swanson, MPH, Patricia Vranesich, BSN, and Walter Orenstein, MD

 See also Galea and Vaughan, p. 1932.

Objectives. To assess the impact of the COVID-19 pandemic on immunization services across the life course.

Methods. In this retrospective study, we used Michigan immunization registry data from 2018 through September 2020 to assess the number of vaccine doses administered, number of sites providing immunization services to the Vaccines for Children population, provider location types that administer adult vaccines, and vaccination coverage for children.

Results. Of 12 004 384 individual vaccine doses assessed, 48.6%, 15.6%, and 35.8% were administered to children (aged 0–8 years), adolescents (aged 9–18 years), and adults (aged 19–105 years), respectively. Doses administered overall decreased beginning in February 2020, with peak declines observed in April 2020 (63.3%). Overall decreases in adult doses were observed in all settings except obstetrics and gynecology provider offices and pharmacies. Local health departments reported a 66.4% decrease in doses reported. For children, the total number of sites administering pediatric vaccines decreased while childhood vaccination coverage decreased 4.4% overall and 5.8% in Medicaid-enrolled children.

Conclusions. The critical challenge is to return to prepandemic levels of vaccine doses administered as well as to catch up individuals for vaccinations missed. (*Am J Public Health.* 2021;111(11):2027–2035. <https://doi.org/10.2105/AJPH.2021.306474>)

The pandemic spread of severe acute respiratory syndrome coronavirus 2 (SARS-CoV-2), the virus that causes COVID-19, exploded onto the global stage in early December 2019 in Wuhan, China.¹ The first confirmed case of COVID-19 in the United States was reported on January 20, 2020,² and community transmission was detected in February 2020. By mid-March, all 50 states, the District of Columbia, and all 4 US territories had reported cases of COVID-19.³ Michigan reported its first case on March 10, 2020, and, soon thereafter, on March 13, 2020, the United States declared a

national state of emergency to control the pandemic spread of the virus.⁴ As of April 27, 2021, more than 918 000 cases and 17 400 deaths have been recorded in Michigan, and more than 31.9 million cases and more than 569 700 deaths have been reported across the United States.³

Public health response measures were implemented across the nation to mitigate the pandemic, centering on social distancing and quarantine policies, including shelter-in-place and stay-at-home orders. Michigan declared a state of emergency⁵ and implemented executive orders to suspend in-person

operations that were not critical to sustain or maintain life, disrupting access to routine health services.^{6,7} Michigan reopened using a tiered approach on June 22, 2021. Public schools, kindergarten through grade 12, also opened for the 2020–2021 academic year through a combination of in-person, online, and hybrid learning.

Core preventive services, including immunizations, were dramatically interrupted resulting in declines in vaccination coverage leaving communities at risk for vaccine-preventable diseases (VPDs) and associated complications.^{8,9} With interruptions of immunization

services, concerns about outbreaks of disease, particularly measles, which requires an estimated immunity level of at least 92% to 94% to reach the herd immunity threshold and prevent outbreaks, were of primary concern.¹⁰ Reductions in measles coverage and coverage of other routinely recommended vaccines (e.g., diphtheria, mumps, pertussis, polio, varicella, and rubella) threaten herd immunity, particularly in areas where unvaccinated persons are in close proximity and introductions of cases through travel are more likely. The critical challenge for Michigan and other states across the nation is to ensure a return to prepandemic levels of doses administered while ensuring catch-up doses are administered to individuals who have fallen behind on complying with the Advisory Committee on Immunization Practices (ACIP) routine immunization schedules.¹¹

To do this, states rely on programs like the Vaccines for Children (VFC) program, a federal safety net that seeks to ensure that all children have access to vaccines without financial barriers. Children eligible for the VFC program include those who are Medicaid-eligible, uninsured, and American Indian or Alaska Native. In addition, children with health insurance that does not cover vaccination (underinsured) can get free vaccines at federally qualified health centers.¹² Furthermore, all states including Michigan utilize pharmacists and pharmacies as vaccinators and sites of service. Pharmacies offer the convenience of extended hours outside of those for a traditional medical clinic, including holidays, expanding access points for routine vaccination services, particularly for adults who value the convenience pharmacies offer.¹³ In this retrospective

study, we used the Michigan Care Improvement Registry (MCIR; the state's immunization information system [IIS]) and assessed vaccine doses administered across the life course during the COVID-19 pandemic. We also assessed routine vaccination coverage for children aged 19 to 35 months.¹⁴

METHODS

The study period for this retrospective analysis of doses reported to MCIR for individuals aged 0 through 105 years was from January 1, 2018, through September 30, 2020. We compared monthly dose administrations from January through September 2020 with January through September 2018–2019 averages. We chose a 2-year average as a representative baseline to account for data fluctuations that may occur in a given year, beginning in 2018, because administration of the 2-dose herpes zoster vaccine began in earnest in 2018, after the November 2017 ACIP recommendation, as supply constraints limited initial vaccinations.^{15–17} The analysis of vaccination coverage for children aged 19 to 35 months included coverage in September 30, 2020, compared with coverage in September 30, 2019, as point estimates.

Data Source

As of September 2020, MCIR contained more than 153 million provider-verified immunization records for more than 11 million individuals, including 2 693 310 children and adolescents aged younger than 19 years and 8 772 125 adults. Duplicate doses, doses from individuals who opted out of the registry, nonvaccine products (e.g., tuberculin skin tests and immunoglobulins), and other doses (e.g., historical

doses, not administered by the reporting MCIR provider—rather, reported based on vaccine documentation) were excluded. Doses were limited to 1 vaccine type (based on vaccine administration code) per day per person (e.g., only 1 hepatitis B included per person per day). Influenza vaccinations were excluded because of the seasonality of vaccine administration and because the study period ended before the start of the 2020–2021 influenza season. Travel and counterbioterrorism vaccines were included; however, they accounted for less than 1% of vaccine doses. Immunization providers are required to report vaccinations administered to persons less than 20 years of age to the MCIR within 72 hours, unless the parent, guardian, or person in loco parentis of the child who received the vaccine objects by written notice.¹⁸ According to annual data reported to the Centers for Disease Control and Prevention, a majority of doses are reported electronically within 24 hours.

Study Population

The following study cohorts were created to assess doses administered in 2020 compared with the monthly average of the 2 previous years (2018–2019) for the months of January to September: young children (aged 0 to < 2 years), children (aged 2 to 8 years), adolescents (aged 9 to 18 years), and adults (aged 19 to 105 years). First, pediatric and adolescent doses administered were reported by programmatic variables: by VFC Program status and whether the doses were administered at a VFC site. Second, doses administered were reported by ACIP-recommended age groups, for adolescents and adults respectively: human papillomavirus;

tetanus, diphtheria, and acellular pertussis (Tdap) or tetanus–diphtheria (Td); and meningococcal serogroup A, C, W, Y (MenACWY) vaccines for adolescents aged 11 through 18 years; and Tdap or Td (adults aged 19 years and older), herpes zoster (for adults aged 50 years and older), and pneumococcal polysaccharide vaccine–23 valent (PPSV23, for adults aged 65 years and older) for adults. Lastly, doses administered for adults were reported by 10 MCIR facility types (i.e., site of service) that most frequently report administering adult vaccinations.

Vaccination coverage (up-to-date status) for children aged 19 to 35 months was also reported. This includes the percentage of children (19–35 months) fully immunized with 4 or more doses of diphtheria, tetanus, and acellular pertussis (DTaP); 3 or more doses of polio; 1 or more dose of measles, mumps, and rubella; 3 or more doses of *Haemophilus influenzae* type b (3 or 4 doses depending on vaccine product), 3 or more doses of hepatitis B, 1 or more dose of varicella, and 4 or more doses of pneumococcal conjugate.¹⁴ We also assessed the change in vaccination coverage in Medicaid-enrolled children and the number of children enrolled in Medicaid during the same 1-year time period.

We analyzed data by using SAS version 9.4 (SAS Institute Inc, Cary, NC).

RESULTS

Of the 12 004 384 provider-verified doses in the study sample, 48.6% were administered to children aged 0 to 8 years, 15.6% to adolescents aged 9 to 18 years, and 35.8% to adults aged 19 years and older. Total doses administered in each age group increased in January 2020 compared with the

average during the same time period in 2018 and 2019, with the largest increase in adults aged 19 through 105 years (32.3%), primarily driven by increased administrations of a newly licensed¹⁵ and ACIP-recommended¹⁶ 2-dose herpes zoster vaccine (Figure 1). The greatest decreases across all age groups were observed in April 2020, with the largest percentage decrease of 85.6% observed among adolescents (2018–2019 average = 46 295; 2020 = 6678), followed by a 82.7% decrease in children aged 2 through 8 years (2018–2019 average = 37 890; 2020 = 6571), a 82.2% decrease in adults (2018–2019 average = 128 383; 2020 = 22 865), and a 34.9% decrease in children aged younger than 2 years (2018–2019 average = 147 456; 2020 = 95 998).

Only 1 age group and time point showed a percentage increase during the pandemic. Specifically, in June 2020, children aged younger than 2 years received 5.7% more vaccines compared with the average in June 2018 and 2019. However, the increase was not sustained, as doses administered in July through September declined, compared with July through September 2018–2019 averages, by 3.5%, 8.6%, and 2.8%, respectively (Figure 1).

Vaccination Sites of Services for Children

The total number of sites reporting vaccines to the MCIR for children aged 0 through 18 years declined 36.2% from March to April 2020; the number of sites reporting VFC and non-VFC doses declined from 1123 and 1634 sites to 718 and 1040 sites, respectively, between March and April 2020. The number of sites reporting doses has

somewhat recovered; however, the total number of sites reporting doses remain below 2018–2019 averages (9384 in 2020 relative to 10 738 in 2018–2019 for VFC sites and 14 899 in 2020 relative to 17 825 in 2018–2019 for non-VFC sites). In addition, the number of VFC doses reported (775 736) was lower than non-VFC doses (1 041 055) reported during the study period.

Doses Administered to Adolescents

Total doses of routinely recommended adolescent vaccines administered declined, with the greatest decrease occurring in 11 through 18-year-old adolescents in April 2020 (Figure 2). Decreases were observed in all routinely recommended adolescent vaccines relative to the same month average during 2018 and 2019 with the peak decline in April: 85.8% decrease in human papillomavirus (from 14 388 to 2043), 86.2% decrease in MenACWY (from 121 156 to 1675), and an 82.6% decrease in Tdap (from 7337 to 1273). Only MenACWY is required for school entry at this age.¹⁹

Doses Administered to Adults

Dramatic decreases were also observed in the routinely recommended adult vaccines evaluated in April: 89.6% decrease in PPSV23 (from 6905 to 716), 85.5% decrease in herpes zoster (from 22 262 to 3229), and 67.5% decrease in Tdap or Td (from 38 802 to 12 616; Figure 3). Doses of herpes zoster and PPSV23 administered during June through September 2020 exceeded doses administered in the same time period (June through September) 2018–2019 averaged.

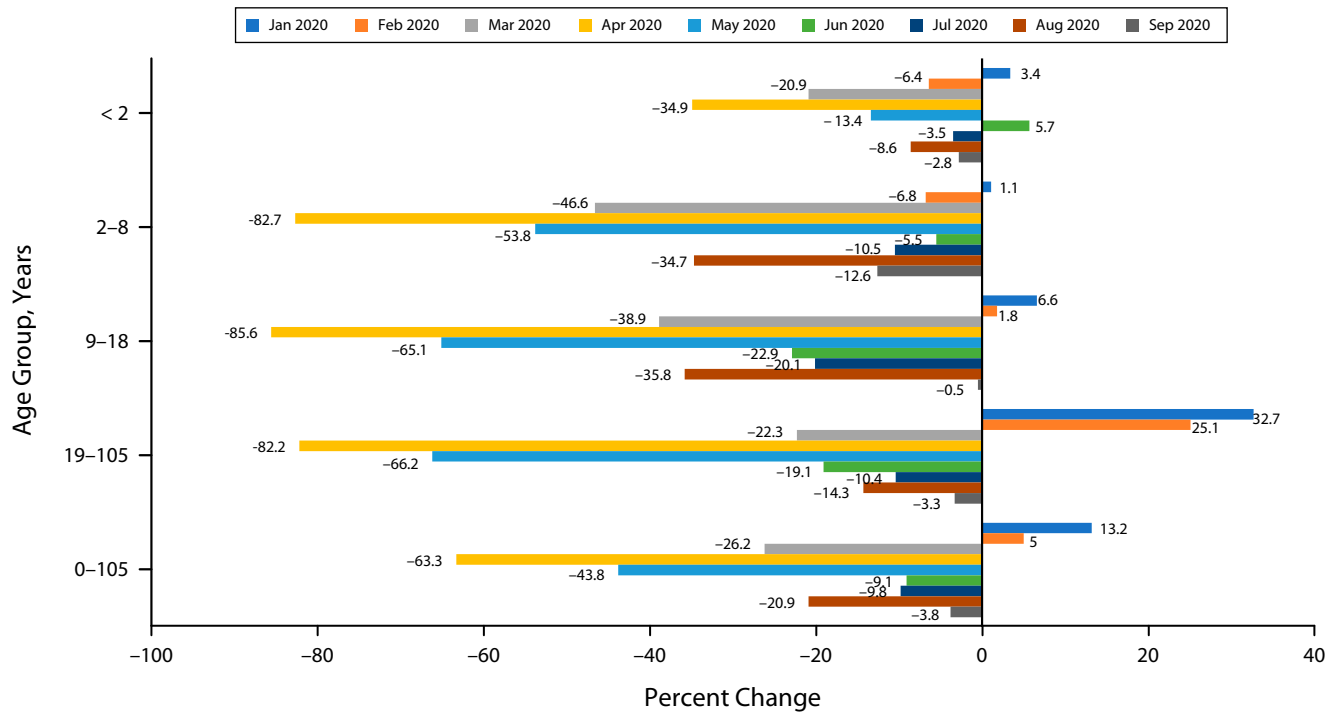
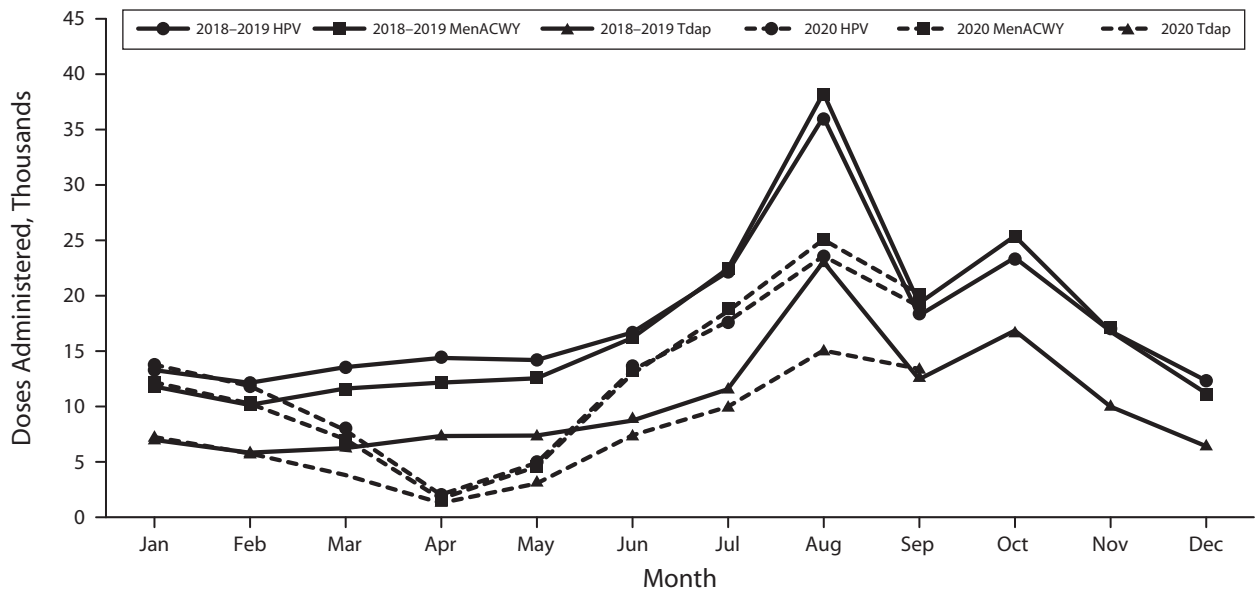


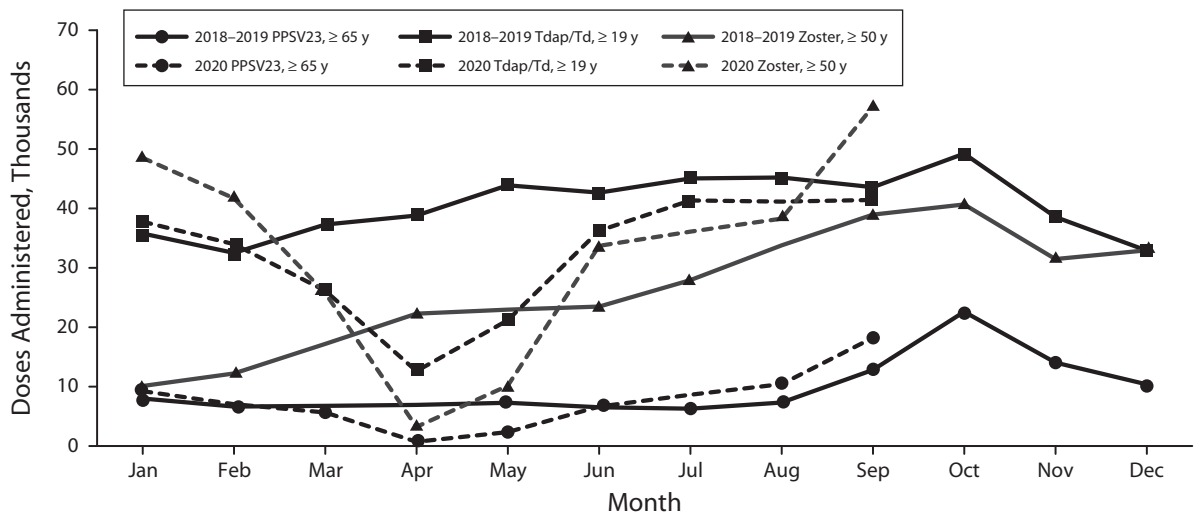
FIGURE 1— Percent Change in Vaccine Doses Administered, by Age Group and Month: Michigan Care Improvement Registry, January–September 2020 vs January–September 2018–2019 Average



	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
2018–19 HPV	13 290	12 134	13 546	14 388	14 198	16 666	22 139	36 121	18 226	23 418	16 798	12 332
2018–19 MenACWY	11 803	10 138	11 608	12 156	12 534	16 190	22 424	38 291	19 322	25 374	16 759	11 188
2018–19 Tdap	6 979	5 834	6 239	7 337	7 365	8 729	11 556	23 054	12 458	16 847	9 989	6 396
2020 HPV	13 791	11 843	7 871	2 043	4 914	13 284	17 670	23 569	19 029			
2020 MenACWY	12 193	10 259	7 060	1 675	4 562	12 948	18 573	25 051	20 151			
2020 Tdap	7 232	5 759	3 802	1 273	3 049	7 388	9 974	15 043	13 395			

FIGURE 2— Doses Administered of Select Vaccine Types, by Month, to Adolescents Aged 11 Through 18 Years: Michigan Care Improvement Registry, January–September 2020 vs January–September 2018–2019 Average

Note. HPV = human papillomavirus; MenACWY = meningococcal serogroup A, C, W, Y; Tdap = tetanus, diphtheria, and acellular pertussis.



	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
2018–2019 PPSV23, ≥ 65 y	8 008	6 643	6 742	6 905	7 267	6 518	6 296	7 318	12 810	22 608	14 035	10 392
2018–2019 Tdap/Td, ≥ 19 y	35 777	32 513	37 274	38 802	43 893	42 642	45 033	45 198	43 556	49 235	38 552	32 924
2018–2019 Zoster, ≥ 50 y	10 098	12 216	17 187	22 262	22 970	23 463	27 945	33 792	38 961	40 680	31 471	32 932
2020 PPSV23, ≥ 65 y	9 234	7 043	5 647	716	2 323	6 652	8 634	10 440	18 233			
2020 Tdap/Td, ≥ 19 y	37 779	33 971	26 136	12 616	21 154	36 251	41 359	41 165	41 430			
2020 Zoster, ≥ 50 y	48 499	41 707	25 746	3 229	10 124	33 690	36 076	38 282	57 147			

FIGURE 3— Doses of Routinely Recommended Adult Vaccines Administered, by Month: Michigan Care Improvement Registry, January–September 2020 vs January–September 2018–2019 Average

Note. PPSV23 = pneumococcal polysaccharide vaccine–23 valent; Td = tetanus–diphtheria; Tdap = tetanus, diphtheria, and acellular pertussis.

The sites of service where adults received vaccines did not change during the pandemic relative to before the pandemic (Table 1). The largest proportion of doses administered to adults aged 50 through 64 years and adults aged 65 years and older were reported by family practice and pharmacy locations (44.0% and 55.8%, respectively); doses administered to adults aged 19 through 49 years (37.8%) were most often reported by family practice sites and hospitals (37.8%). The greatest decreases were observed in doses reported by local health departments for adults aged 19 through 49 years (70.3%) and adults aged 50 through 64 years (66.1%). Obstetrics and gynecology provider sites and pharmacies were the only locations that reported percentage increases in doses administered to adults during 2020 (11.1% and

0.4%, respectively). Any remaining types were defined as “other” (Table 1).

Childhood Vaccination Coverage

Recommended vaccination coverage for children aged 19 through 35 months for the complete doses of the combined 7-vaccine series to prevent 13 diseases was 70.3% as of September 30, 2020, a decrease of 4.4% from September 30, 2019. Coverage for completion of individual vaccines also declined; specifically, coverage with DTaP was 72.9% (a 4.4% decrease from 2019). Vaccination coverage and respective declines for the completion of individual vaccines was 79.3% for pneumococcal conjugate (3.9%); 82.4% for measles, mumps, and rubella (3.6%); 81.4% for varicella (3.4%); 83.9%

for *Haemophilus influenzae* type b (3.1%); 85.2% for hepatitis B (2.2%); and 84.7% for polio (1.6%). From September 30, 2019, to September 30, 2020, the number of Medicaid-enrolled children aged 19 through 35 months in MCIR increased 4.3%, from 71 798 to 74 837, while vaccination coverage for the complete series of the combined 7-vaccine series of Medicaid-eligible children decreased 5.8%, from 73.1% in 2019 to 67.3% in 2020. Vaccination coverage for the individual vaccines also decreased in Medicaid-eligible children with the largest decrease in DTaP vaccination of 5.6% (69.6% in 2020 vs 75.2% in 2019).

DISCUSSION

As Michigan and the entire country emerged from stringent lock-down

TABLE 1— Adult Vaccine Doses Administered and Percent Change, by Site of Service and Age Group: Michigan Care Improvement Registry, January–September 2020 vs January–September 2018–2019 Average

Facility Site of Service	Adults Aged 19–49 y			Adults Aged 50–64 y			Adults Aged ≥ 65 y			Total, Adults Aged ≥ 19 y		
	2018–2019 (n = 475 156), No. (%)	2020 (n = 315 577), No. (%)	Percent Change	2018–2019 (n = 318 501), No. (%)	2020 (n = 294 370), No. (%)	Percent Change	2018–2019 (n = 391 236), No. (%)	2020 (n = 346 945), No. (%)	Percent Change	2018–2019 (n = 1 184 893), No. (%)	2020 (n = 956 892), No. (%)	Percent Change
	Family practice	188 276 (39.6)	132 499 (27.9)	-29.6	127 439 (26.8)	123 131 (25.9)	-3.4	115 513 (24.3)	94 024 (19.8)	-18.6	431 227 (36.4)	349 654 (36.5)
Pharmacy	32 549 (6.9)	20 365 (4.3)	-37.4	75 505 (15.9)	85 954 (18.1)	13.8	168 267 (35.4)	171 033 (36.0)	1.6	276 320 (23.3)	277 352 (29.0)	0.4
Hospital	56 310 (11.9)	47 020 (9.9)	-16.5	23 259 (4.9)	18 454 (3.9)	-20.7	27 085 (5.7)	21 729 (4.6)	-19.8	106 653 (9.0)	87 203 (9.1)	-18.2
Internal medicine	26 319 (5.5)	17 615 (3.7)	-33.1	28 675 (6.0)	28 425 (6.0)	-0.9	40 605 (8.5)	32 059 (6.7)	-21.0	95 598 (8.1)	78 099 (8.2)	-18.3
Other ^a	25 296 (5.3)	13 825 (2.9)	-45.3	15 127 (3.2)	12 232 (2.6)	-19.1	13 043 (2.7)	11 290 (2.4)	-13.4	53 465 (4.5)	37 347 (3.9)	-30.1
Local health department	60 296 (12.7)	17 934 (3.8)	-70.3	26 097 (5.5)	8 853 (1.9)	-66.1	14 959 (3.1)	7 298 (1.5)	-51.2	101 352 (8.6)	34 085 (3.6)	-66.4
Obstetrician/gynecologist	26 384 (5.6)	29 467 (6.2)	11.7	255 (0.1)	240 (0.1)	-5.9	160 (0.0)	61 (0.0)	-61.8	26 799 (2.3)	29 768 (3.1)	11.1
Pediatrics	17 267 (3.6)	13 267 (2.8)	-23.2	7 371 (1.6)	8 734 (1.8)	18.5	5 451 (1.1)	5 229 (1.1)	-4.1	30 089 (2.5)	27 230 (2.8)	-9.5
Urgent care	10 545 (2.2)	8 509 (1.8)	-19.3	5 087 (1.1)	3 979 (0.8)	-21.8	3 193 (0.7)	2 403 (0.5)	-24.7	18 825 (1.6)	14 891 (1.6)	-20.9
College or university	15 756 (3.3)	8 609 (1.8)	-45.4	3 455 (0.7)	2 220 (0.5)	-35.7	2 170 (0.5)	1 555 (0.3)	-28.3	21 381 (1.8)	12 384 (1.3)	-42.1
Corporate	16 160 (3.4)	6 467 (1.4)	-60.0	6 234 (1.3)	2 148 (0.5)	-65.5	793 (0.2)	264 (0.1)	-66.7	23 187 (2.0)	8 879 (0.9)	-61.7

^aOther includes specialty clinic, travel clinic, correctional clinic, tribal health center, migrant health center, adolescent health center, adult-only clinic, adolescent-only clinic, long-term care, sexually transmitted disease center, family planning clinic, and Veterans Affairs hospital.

measures imposed in spring 2020, adults have resumed work in congregate settings; students, to varying degrees, have physically returned to schools and universities; and individuals have returned to popular community settings, like the gym, bars, and restaurants. In our study, we found dramatic decreases in (1) doses administered across the life course; (2) adult dose administration across the majority of provider settings, particularly in local health departments; and (3) vaccination coverage for the complete doses of the combined 7-vaccine series in children aged 19 to 35 months in 2020 compared with previous data.

As we strive to achieve pre-COVID-19 levels of doses administered for routinely recommended vaccines, it is vital to ensure catch-up vaccination of doses missed throughout the pandemic to stem outbreaks of VPDs (e.g., measles). Decreasing coverage can lead to resurgences of all VPDs. Because measles is the most contagious of the VPDs, measles may be the first such disease to increase.¹⁰ This study also showed a decrease in pediatric immunization provider sites, particularly those dedicated to caring for more vulnerable populations, like Medicaid-insured children; this will complicate the ability to reach pre-pandemic rates quickly with fewer access points for vaccination and, in some cases, fewer appointment options.

To reach prepandemic levels of vaccination coverage and to ensure catch-up vaccination on doses missed since the onset of the pandemic, it is critical to take steps to ensure both adult and pediatric providers identify and target their patients for catch-up doses.^{20–22} The IIS can help with this

effort, as a supplement to provider electronic health records for doses that are not captured by a provider's electronic health records to identify those indicated for vaccination and to contact them to schedule appointments. Additional clinic hours and appointment times can also be a useful strategy to support efforts. At every encounter, providers can ensure that the immunization status of every individual is assessed, vaccines are strongly recommended, and individuals are vaccinated or referred for vaccination.²³ While the rate of decrease in doses administered has slowed since April 2020 (Figure 1), decreases in preventive care during the pandemic are cause for concern. This concern stems from the potential for missed diagnoses that may increase infectious disease susceptibility coupled with health care practices, particularly small, private practices, which continue to struggle to survive in the wake of the COVID-19 pandemic, changing health care patterns of access and utilization.²⁴⁻²⁶

August is often associated with back-to-school vaccinations, and this study confirmed increased doses administered in August 2020 (Figure 2); however, up-to-date vaccination coverage is a more accurate reflection of protection against VPDs, and this will take longer to rebound. While providers and some families appear to be seeking catch-up vaccinations, lower coverage among Medicaid-covered children is concerning.⁸ Increases in Medicaid enrollment from the previous year, most likely attributable to COVID-19-induced economic-related insurance coverage shifts, means more children are shifting from other insurance types onto Medicaid coverage and eligibility for the VFC program, shifting the costs of vaccination onto

public programs. The marginal increase already observed in this study may signal an ongoing potential for increases in Medicaid enrollment, particularly as the long-term economic impact of the pandemic is yet to be fully realized; therefore, monitoring changes in VFC immunization provider sites will be important for ensuring access to vulnerable populations and at-risk communities. Changes in practice patterns may be a result of pay cuts, staff layoffs because of low patient volume, and the consequences of adhering to public health guidelines related to social distancing (e.g., fewer appointments, fewer people in the office).

Dramatic declines in doses administered have yet to translate to dramatic declines in vaccination coverage as the full impact on coverage levels at specific ages (i.e., 19–35 months), as young children aged 2 years and younger who missed doses have yet to age into the 19- through 35-month age cohort assessed by the National Immunization Survey.¹⁴ For example, a child aged 6 months who missed multiple vaccinations throughout the pandemic would not be included in the coverage assessment. The decreases in this study reflect preliminary decreases in coverage, which can be alarming if children are not caught up on recommended vaccines.

We observed age-related differences in adult vaccines as doses administered of PPSV23 and herpes zoster have rebounded to pre-COVID-19 levels. The rebound in doses administered for PPSV23 may reflect a concern for other respiratory pathogens amid SARS-CoV-2 cocirculation. Undoubtedly, resolution of persistent supply issues for herpes zoster and an ACIP-preferential recommendation for a 2-dose zoster vaccine¹⁶ have

contributed to increased doses administered, as the previous zoster vaccine was only a single dose.¹⁷ By contrast, younger adults may not feel as much of a need to seek health care, or they may have difficulties related to appointment availability, causing them to change where they get immunizations, particularly as providers may have decreased appointment “slots.” Some special populations such as pregnant women were prioritized as a priority population for appointments (Table 1). Even as providers consider different adult populations and pandemic-associated changes in vaccination-seeking behaviors, it is important that even those not offering, or no longer offering, vaccinations continue to promote the importance of vaccination.²³

Pharmacies have increasingly become a vital vaccination resource in communities. Michigan has observed an increase in the number of pharmacy sites and the number of doses reported to the MCIR. There has been an increase in electronic reporting by pharmacies, which play a critical role in administering vaccinations, particularly to older adults (Table 1).^{23,27,28} Ensuring individuals are vaccinated, particularly as other respiratory pathogens (e.g., influenza, *Streptococcus pneumoniae*), are circulating during the COVID-19 pandemic, has been a priority.²⁹ While this study period ended before influenza vaccinations began in earnest, efforts to vaccinate for routinely recommended vaccines are important, particularly in adults aged 65 years and older, as this cohort is a vulnerable population accounting for a disproportionate number of seasonal influenza-related deaths, an estimated 70% to 85%, and a high proportion of influenza-related hospitalizations.³⁰ This same population

accounts for 8 of every 10 confirmed COVID-19–reported deaths in the United States.^{31,32}

Limitations

Vaccination coverage is the traditional metric and gold standard used to assess vaccine utilization, a population's level of protection from disease, and immunization program performance; however, doses administered and reported to an IIS represent timely and available proxy measures. One well-established limitation in the use of IIS data, particularly in the adolescent and adult population, is denominator inflation.^{33,34} Denominator inflation occurs when IIS client totals exceed population estimates, which can skew population coverage estimates, for example, because of multiple records for 1 individual. This inflation may underestimate coverage rates. Overestimates of coverage estimates occur when a population estimate is used for the denominator. We attempted to account for this by conducting manual checks to deduplicate data and limiting coverage calculations to childhood vaccinations.

Another limitation is that adult immunization reporting is not required in Michigan; therefore, completeness of reporting for adults is lower than for children. It is possible that some of the changes noted in the number of reported adult immunizations in 2020 were attributable to reductions in reporting rather than changes in administration. In responding to the pandemic, we suspect that providers may have had less time because of competing priorities to report all administered doses to the registry.

A third limitation is that the classification of MCIR facility types is performed

by MCIR regional staff, who are best positioned to know the populations served by the facilities in their regions; however, these classifications are not routinely verified and updated after enrollment.

Lastly, at the time of this analysis, race and ethnicity data, which can be useful in understanding the health disparities and health care utilization patterns of care in vulnerable communities, were incomplete across the life course.

Public Health Implications

After a 2020 summer lull in daily COVID-19 case counts, cases began to rise exponentially, with unprecedented record metrics (e.g., hospitalizations) associated with the pandemic in November 2020.³ Diligence in monitoring vaccination rates and provider site availability will be critical to returning to prepandemic levels of coverage as well as catching up cohorts of individuals for vaccinations. It will also be critical for all providers to be astute in considering VPD diagnoses in their diagnostic assessments moving forward. As vaccination coverage has dropped, susceptible populations are accumulating, which could lead to outbreaks of VPDs not seen in recent times. Given that measles is the most contagious of the VPDs, outbreaks of measles may be the first evidence of resurgences in VPDs. The susceptibility gap for all of the VPDs needs to be closed as soon as possible.

Since May 2021, COVID-19 vaccination programs are now widely available across all jurisdictions in the United States. As a result of increased capability to report race and ethnicity data for adults, as required for COVID-19 vaccine, Michigan and other jurisdictions now have enhanced tools to address

under- and unvaccinated populations. As society shifts to a new normal, recalibrating to a world where SARS-CoV-2 is endemic, COVID-19 vaccines will certainly transition onto the routine immunization schedule in some form. It is critical to ensure the immunization delivery system supports timely, accessible, and reliable access to routinely recommended vaccines across the nation, sustaining historical high coverage in children and strengthening increasing coverage for adolescents and adults. *AJPH*

ABOUT THE AUTHORS

Angela K. Shen is a visiting scientist with the Children's Hospital of Philadelphia, Philadelphia, PA. Crista A. Bramer, Lynsey M. Kimmins, Robert Swanson, and Pat Vranesich are with the Michigan Department of Health and Human Services, Division of Immunization, Lansing. Walt Orenstein is with Emory University School of Medicine, Atlanta, GA.

Note. The views here do not represent the official views of the Michigan Department of Health and Human Services.

CORRESPONDENCE

Correspondence should be sent to Angela K. Shen, ScD, MPH, at Children's Hospital of Philadelphia, Vaccine Education Center, 3615 Civic Center Boulevard, Suite 1202 ARC, Philadelphia, PA 19104 (e-mail: shenak@chop.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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A. K. Shen, C. A. Bramer, and L. M. Kimmins conceptualized and designed the study, and analyzed and interpreted the data for the study. A. K. Shen wrote the first draft. All authors were equally responsible for critical review and approval of the final article.

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The authors have no conflicts of interest to disclose.

HUMAN PARTICIPANT PROTECTION

This study was considered exempt as a program evaluation of a public health activity by the Michigan Department of Health and Human Services institutional review board.

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Workplace Leave and Breastfeeding Duration Among Postpartum Women, 2016–2018

Katherine Kortzmit, PhD, MPH, Rui Li, PhD, Shanna Cox, MSPH, Carrie K. Shapiro-Mendoza, PhD, MPH, Cria G. Perrine, PhD, Denise V. D'Angelo, MPH, Wanda D. Barfield, MD, MPH, Holly B. Shulman, MA, Craig F. Garfield, MD, MAPP, and Lee Warner, PhD, MPH

Objectives. To examine associations of workplace leave length with breastfeeding initiation and continuation at 1, 2, and 3 months.

Methods. We analyzed 2016 to 2018 data for 10 sites in the United States from the Pregnancy Risk Assessment Monitoring System, a site-specific, population-based surveillance system that samples women with a recent live birth 2 to 6 months after birth. Using multivariable logistic regression, we examined associations of leave length (< 3 vs ≥ 3 months) with breastfeeding outcomes.

Results. Among 12 301 postpartum women who planned to or had returned to the job they had during pregnancy, 42.1% reported taking unpaid leave, 37.5% reported paid leave, 18.2% reported both unpaid and paid leave, and 2.2% reported no leave. Approximately two thirds (66.2%) of women reported taking less than 3 months of leave. Although 91.2% of women initiated breastfeeding, 81.2%, 72.1%, and 65.3% of women continued breastfeeding at 1, 2, and 3 months, respectively. Shorter leave length (< 3 months), whether paid or unpaid, was associated with lower prevalence of breastfeeding at 2 and 3 months compared with 3 or more months of leave.

Conclusions. Women with less than 3 months of leave reported shorter breastfeeding duration than did women with 3 or more months of leave. (*Am J Public Health.* 2021;111(11):2036–2045. <https://doi.org/10.2105/AJPH.2021.306484>)

Breast milk is recognized globally as the ideal form of nutrition for most infants for optimal growth and development.^{1,2} Improving US breastfeeding rates is a public health priority.^{3,4} In the United States, the American Academy of Pediatrics recommends that mothers exclusively breastfeed for about 6 months and continue breastfeeding as complementary foods are introduced through the infant's first birthday.¹ Although most mothers initiate breastfeeding,⁵ many face multiple barriers to continuing.^{4,6,7} Specifically, mothers employed outside the home

face unique challenges, including separation from their infants when returning to work and inadequate time or space to express milk at work, which can lead to early cessation of breastfeeding.^{6,7} Over the past half century, the number of first-time mothers participating in the workforce has increased, with the percentage who worked during pregnancy increasing from 44% in 1961 to 1965 to 66% in 2006 to 2008.⁸ In 2018, nearly two thirds of women who had a live birth in the past year were in the workforce in the United States.⁹

The US Surgeon General's 2011 *Call to Action to Support Breastfeeding* outlined actions employers could take to support employees who breastfeed, including establishing paid maternity leave and lactation support programs.⁴ Policies that support maternal leave and breastfeeding for women in the workplace include the Family Medical Leave Act (FMLA),¹⁰ the Patient Protection and Affordable Care Act,¹¹ paid family leave policies enacted or passed in 8 states and the District of Columbia,¹² and, for federal employees, the Federal Employee Paid Leave Act (effective

October 2020).¹³ In addition to leave policies at the state and federal levels, several large organizations offer their employees paid family leave.¹⁴ Although the FMLA (up to 12 weeks of unpaid leave)¹⁰ and paid family leave policies¹²⁻¹⁴ provide opportunities for mothers to take leave after delivery, many cannot afford to take unpaid leave, are not covered by the policies, or do not meet the eligibility criteria (e.g., length of time employed, number of hours worked) to participate.⁴ Limited access to leave means many women are also returning to the workforce soon after giving birth.⁸

Research examining data before the Surgeon General's *Call to Action*¹⁵⁻¹⁹ and, more recently, state-specific examinations on the effects of paid leave^{20,21} and small-scale studies on specific populations (e.g., military)²² have demonstrated that women who are able to remain on leave longer are also more likely to continue breastfeeding. Population-based analyses that consider both paid and unpaid leave are lacking. We compared the prevalence of breastfeeding initiation and any breastfeeding at 1, 2, and 3 months by length of leave taken, both paid and unpaid, among a large representative sample of recently postpartum women who gave birth during January 2016 to December 2018.

METHODS

We derived data from the Pregnancy Risk Assessment Monitoring System (PRAMS), a multisite, population-based surveillance system. During the years examined, PRAMS collected data on maternal attitudes, behaviors, and experiences before, during, and shortly after pregnancy using a standardized questionnaire and protocol from 47 states, the District of Columbia, New York City

(NYC), and Puerto Rico (hereafter described as "sites"). PRAMS sites selected a stratified random sample of women with a recent live birth from site birth certificate files 2 to 6 months after birth. Sampled women were mailed a self-administered survey. Following non-response to 3 mailed surveys, PRAMS sites initiated telephone follow-up (up to 15 calls). Each site's PRAMS survey included a mandatory "core" questionnaire, and each site had the option to include additional "standard" questions from a library of optional question modules that expanded on or addressed different topics not captured by core questions. We analyzed PRAMS 2016 to 2018 data from 10 sites (Massachusetts, Maryland, Minnesota, Missouri, North Carolina, New Hampshire, New York State, Vermont, Wisconsin, and NYC) that included standard workplace leave-related questions on their site-specific survey and achieved a weighted response rate of 55% or greater for at least 1 year during the study period. The Centers for Disease Control and Prevention weights PRAMS data for sampling design, noncoverage, and nonresponse to be representative of each site's live birth population. PRAMS sites mailed sampled women a written informed consent with the survey. For those who completed the survey during telephone follow-up, informed consent was completed before starting the survey. Further detail on PRAMS methodology has been described elsewhere.²³

Measures

Type of leave and leave length. Women who were employed during pregnancy and had returned (or planned to return) to the same job they had during pregnancy after giving birth reported

on the type of leave they took and the length of leave they had taken or planned to take. Respondents were asked, "Did you take leave from work after your new baby was born? (check all that apply)" and asked to respond from the following options: "I took paid leave from my job," "I took unpaid leave from my job," and "I did not take any leave." NYC and Missouri included site-specific responses on type of leave. We coded the NYC response option "I took leave and used temporary disability insurance" as paid leave. In NYC, 29.7% of women who took leave reported receiving temporary disability insurance. We coded the Missouri response option "Family Medical Leave (paid or unpaid)" as unpaid leave based on Missouri Family Medical Leave laws.²⁴

We coded type of leave into 4 categories: (1) "paid leave only," (2) "unpaid only," (3) "both paid and unpaid leave," and (4) "no leave." Women who reported taking any leave were also asked, "How many weeks or months of leave, in total, did you take or will you take?" We categorized leave length as less than 3 months (≤ 12 weeks; this included women reporting no leave) and 3 or more months (≥ 13 weeks) of leave. We selected this categorization, as women might have qualified for up to 12 weeks of leave under the FMLA,¹⁰ and this categorization has been used previously in research assessing breastfeeding outcomes.²⁵ We also examined the following 3-level categorization of leave length—0 to 5 weeks, 6 to 12 weeks, and 13 or more weeks of leave—as this categorization has also been used in previous research on breastfeeding outcomes.²⁵ We did not find differences in breastfeeding outcomes between women with 0 to 5 and 6 to 12 weeks of leave (Table A

[available as a supplement to the online version of this article at <http://www.ajph.org>]; therefore, we report on leave categorized as less than 3 versus 3 or more months of leave. To describe whether any leave was paid, we also created a dichotomous indicator distinguishing “no paid leave” (this included women reporting no leave and unpaid leave only) and “any paid leave” (this included women reporting paid leave only or both paid and unpaid leave).

Breastfeeding initiation and any breastfeeding at 1, 2, and 3 months. To measure breastfeeding initiation, respondents were asked, “Did you ever breastfeed or pump breast milk to feed your new baby, even for a short period of time?” We used 2 questions to measure breastfeeding duration: (1) women who ever breastfed were asked, “Are you currently breastfeeding or feeding pumped milk to your new baby?” and (2) women who had stopped breastfeeding when they completed the PRAMS survey were asked, “How many weeks or months did you breastfeed or pump milk to feed your baby?” Women whose infant was deceased or not living with them when they completed the survey were instructed to skip breastfeeding-related questions.

We created the following 4 dichotomous yes–no indicators for breastfeeding: (1) breastfeeding initiation, (2) any breastfeeding at 1 month (4 weeks), (3) any breastfeeding at 2 months (9 weeks), and (4) any breastfeeding at 3 months (13 weeks).

Statistical Analyses

Our analytic sample excluded women whose infants were deceased or not living with them at time of survey completion, women who did not work for pay

during pregnancy, those who were not returning to the same job they had during pregnancy, those who were aged 17 years or younger at time of delivery (because of federal and state-level age restrictions on work hours),²⁶ and those who were missing data on covariates or leave type and length. We also excluded women who reported inconsistent information between leave type and duration from the analysis. After excluding those with missing or discordant data between leave type and length (3.2% and 4.3%, respectively), missing data on breastfeeding initiation and duration (0.2% and 1.2%, respectively) and covariates (5.5%), our final analytic sample included 12 301 (weighted $n = 718\,139$) women who had worked during pregnancy and had returned or planned to return to the same job after giving birth and for whom leave length and breastfeeding were known.

We performed descriptive statistics (the χ^2 test and 95% confidence intervals [CIs]) to assess leave length (< 3 and ≥ 3 months) overall and separately for selected characteristics and by PRAMS site. We identified selected characteristics a priori based on measures that have been associated with leave or breastfeeding outcomes.^{15,19} Data for these characteristics came from birth certificate data available in the PRAMS data set and from PRAMS survey data. Selected characteristics from birth certificate data included maternal race and Hispanic origin (Hispanic, non-Hispanic White, non-Hispanic Black, and non-Hispanic other), age (18 – 24, 25 – 34, and ≥ 35 years), education (\leq high school diploma or general equivalency diploma, some college or associate’s degree, and bachelor’s degree or higher), marital status (married and

unmarried), parity (primiparous and multiparous), and infant gestational age (preterm: < 37 weeks; term: ≥ 37 weeks). Federal poverty level ($\leq 100\%$, > 100%–200%, and > 200%) was available from PRAMS survey data. We also examined leave length by type of leave (no paid leave and any paid leave).

We constructed 4 separate models to describe the associations of each breastfeeding outcome (breastfeeding initiation and breastfeeding at 1, 2, and 3 months) with leave length (< 3 and ≥ 3 months of leave). We calculated the model-based prevalence estimate for each breastfeeding outcome with predicted marginal means and then estimated both unadjusted prevalence ratios (PRs) and adjusted prevalence ratios (APRs) and their associated 95% CIs for leave length. Each model adjusted for all previously mentioned characteristics, timing of survey completion (< 6 vs ≥ 6 months after giving birth), and PRAMS site.

Because previous research has shown that the association of leave length with breastfeeding outcomes varies by select characteristics,²¹ we also examined interactions. For each of the selected characteristics previously mentioned, we constructed a separate model, which included an interaction term between the respective characteristic being examined and leave length. If there was a significant interaction ($P < .01$ based on the F-test for 2-way interaction), we stratified results by the respective characteristic. We also constructed a model to examine the interaction term between leave length and paid leave, and we report these stratum-specific results. Each model contained only 1 interaction term and adjusted for all other selected characteristics.

For all analyses examining breastfeeding at 3 months, we restricted the sample to those who completed the PRAMS survey 3 or more months after delivery ($n = 10\,031$). We performed sensitivity analyses on other breastfeeding outcomes among this restricted sample to assess the robustness of our results. We conducted all analyses with SAS version 9.4 (SAS Institute, Cary, NC) and SAS-callable SUDAAN version 11.0.1 (RTI

International, Research Triangle Park, NC) using weighted data to account for the complex sampling design of PRAMS.

RESULTS

In our study, there were 12 301 women with a recent live birth who had returned or planned to return to the same job they had during pregnancy after giving birth. Among these women,

97.8% reported taking leave (42.1% reported taking only unpaid leave, 37.5% reported only paid leave, and 18.2% reported both unpaid and paid leave). Women who reported taking or planning to take leave had a mean of 12 weeks of leave (median = 11 weeks), with 66.2% of women reporting less than 3 months of leave and 33.8% reporting 3 or more months of leave (Table 1). By site, the prevalence of 3 or more months of leave ranged from

TABLE 1— Prevalence of Leave Length After Delivery Among Women With a Recent Live Birth by Selected Characteristics: Pregnancy Risk Assessment Monitoring System, 10 US Sites, 2016–2018

Characteristics ^a	Total No. ^b	Leave Length		
		< 3 Mo, % (95% CI) ^c	≥ 3 Mo, % (95% CI) ^c	χ^2 P
Total	12 301	66.2 (65.1, 67.3)	33.8 (32.7, 34.9)	
Type of leave ^d				< .001
No paid leave	5 670	70.4 (68.8, 72.0)	29.6 (28.0, 31.2)	
Any paid leave	6 631	62.8 (61.3, 64.3)	37.2 (35.7, 38.7)	
Maternal race and Hispanic origin				< .001
Hispanic	1 276	56.8 (53.0, 60.6)	43.2 (39.4, 47.0)	
Non-Hispanic White	7 303	69.2 (67.9, 70.5)	30.8 (29.5, 32.1)	
Non-Hispanic Black	2 177	61.2 (58.1, 64.2)	38.8 (35.8, 41.9)	
Non-Hispanic other ^e	1 545	61.5 (58.0, 64.9)	38.5 (35.1, 42.0)	
Maternal age, y				< .001
18–24	1 506	78.0 (74.9, 80.8)	22.0 (19.2, 25.1)	
25–34	7 800	66.9 (65.5, 68.3)	33.1 (31.7, 34.5)	
≥ 35	2 995	57.8 (55.4, 60.1)	42.2 (39.9, 44.6)	
Maternal education				< .001
≤ high school diploma or GED	2 199	71.2 (68.4, 73.9)	28.8 (26.1, 31.6)	
Some college or associate's degree	3 337	73.2 (71.1, 75.1)	26.8 (24.9, 28.9)	
Bachelor's or higher degree	6 765	61.2 (59.7, 62.7)	38.8 (37.3, 40.3)	
Marital status				< .001
Married	8 522	64.5 (63.2, 65.8)	35.5 (34.2, 36.8)	
Unmarried	3 779	70.1 (67.9, 72.2)	29.9 (27.8, 32.1)	
Household income by federal poverty level, %				< .001
≤ 100	2 147	72.0 (69.1, 74.6)	28.0 (25.4, 30.9)	
> 100–200	2 333	73.7 (71.2, 76.0)	26.3 (24.0, 28.8)	
> 200	7 821	62.6 (61.2, 64.0)	37.4 (36.0, 38.8)	
Parity				.001
Primiparous	5 524	63.9 (62.2, 65.6)	36.1 (34.4, 37.8)	
Multiparous	6 777	67.9 (66.4, 69.4)	32.1 (30.6, 33.6)	

Continued

TABLE 1— Continued

Characteristics ^a	Total No. ^b	Leave Length		
		< 3 Mo, % (95% CI) ^c	≥ 3 Mo, % (95% CI) ^c	χ^2 P
Gestational age at delivery, wk				< .001
Preterm, < 37	1 992	55.5 (52.0, 59.0)	44.5 (41.0, 48.0)	
Term, ≥ 37	10 309	67.0 (65.9, 68.2)	33.0 (31.8, 34.1)	
Site				< .001
Maryland	1 016	68.6 (65.3, 71.7)	31.4 (28.3, 34.7)	
Massachusetts	2 114	56.6 (54.0, 59.3)	43.4 (40.7, 46.0)	
Minnesota	743	67.8 (64.0, 71.4)	32.2 (28.6, 36.0)	
Missouri	1 503	83.0 (80.8, 85.0)	17.0 (15.0, 19.2)	
North Carolina	411	79.1 (74.3, 83.1)	20.9 (16.9, 25.7)	
New Hampshire	668	69.0 (64.7, 73.0)	31.0 (27.0, 35.3)	
New York State	936	64.7 (60.8, 68.4)	35.3 (31.6, 39.2)	
New York City	1 666	44.3 (41.7, 47.0)	55.7 (53.0, 58.3)	
Vermont	1 593	65.0 (62.6, 67.4)	35.0 (32.6, 37.4)	
Wisconsin	1 651	79.3 (76.6, 81.9)	20.7 (18.1, 23.4)	

Note. CI = confidence interval; GED = general equivalency diploma.

^aType of leave and household income by federal poverty level were obtained from Pregnancy Risk Assessment Monitoring System (PRAMS) survey data. All other characteristics were obtained from birth certificate data available in the PRAMS data set.

^bUnweighted sample size.

^cWeighted % (95% CI).

^dNo paid leave included women reporting no leave and those reporting unpaid leave only. Any paid leave included women reporting paid leave only and those reporting both paid and unpaid leave.

^eIncluded women who self-reported as American Indian, Alaska Native, Asian, Hawaiian, Pacific Islander, or mixed race or other non-White on the birth certificate.

17.0% in Missouri to 55.7% in NYC. Prevalence of taking or planning to take 3 or more months of leave after delivery was highest among women who had any paid leave (37.2%), who were aged 35 years or older (42.2%), who had a bachelor's or higher degree (38.8%), who were married (35.5%), who had a household income level higher than 200% the federal poverty level (37.4%), who were primiparous (36.1%), and whose infant was born preterm (44.5%). Prevalence of taking 3 or more months of leave was also higher among those who were Hispanic (43.2%), non-Hispanic Black (38.8%), or non-Hispanic other (38.5%) than among those who were non-Hispanic White (30.8%). Overall, most (91.2%) women reported initiating

breastfeeding; however, the prevalence of any breastfeeding was lower at both 1 (81.2%) and 2 (72.1%) months. Among those who had completed the PRAMS survey at 3 or more months after birth ($n = 10\,031$), 65.3% reported any breastfeeding at 3 months.

In both unadjusted and adjusted analyses (Table 2), a smaller proportion of women who reported taking or planning to take less than 3 months of leave than those reporting 3 or more months of leave reported ever breastfeeding (90.4% vs 93.2%; APR = 0.97; 95% CI = 0.95, 0.98), breastfeeding at 1 month (79.7% vs 84.5%; APR = 0.94; 95% CI = 0.92, 0.97), and breastfeeding at 2 months (70.1% vs 76.2%; APR = 0.92; 95% CI = 0.89, 0.95). Among those who had completed the PRAMS

survey at 3 or more months, a smaller proportion of women who reported taking or planning to take less than 3 months of leave reported any breastfeeding at 3 months than those with 3 or more months of leave (63.2% vs 69.8%; APR = 0.90; 95% CI = 0.87, 0.94). In sensitivity analyses, when restricting to those who had completed the PRAMS survey at 3 or more months, findings for all breastfeeding outcomes at different periods were consistent with that reported for the full sample (Table B [available as a supplement to the online version of this article at <http://www.ajph.org>]).

For breastfeeding initiation and breastfeeding at 1 month after birth, we observed no interactions between any of the maternal and infant characteristics examined and leave length. For

TABLE 2— Prevalence and Prevalence Ratio of Breastfeeding Initiation and Breastfeeding at 1, 2, and 3 Months by Leave Length Among Women With a Recent Live Birth Who Were Employed During and After Pregnancy: Pregnancy Risk Assessment Monitoring System, 10 US sites, 2016–2018

	Total No. ^a	Unadjusted % (95% CI) ^b	PR (95% CI)	Adjusted % (95% CI) ^{b,c}	APR (95% CI) ^c
Initiated breastfeeding, n = 12 301					
Leave length					
< 3 mo	7866	89.6 (88.6, 90.5)	0.95 (0.94, 0.96)	90.4 (89.5, 91.2)	0.97 (0.95, 0.98)
≥ 3 mo	4435	94.3 (93.3, 95.2)	1 (Ref)	93.2 (92.0, 94.3)	1 (Ref)
Breastfeeding at 1 mo, n = 12 301					
Leave length					
< 3 mo	7866	78.7 (77.4, 80.0)	0.91 (0.89, 0.94)	79.7 (78.5, 80.9)	0.94 (0.92, 0.97)
≥ 3 mo	4435	86.1 (84.6, 87.4)	1 (Ref)	84.5 (82.8, 86.0)	1 (Ref)
Breastfeeding at 2 mo, n = 12 301					
Leave length					
< 3 mo	7866	69.0 (67.6, 70.4)	0.88 (0.86, 0.91)	70.1 (68.7, 71.4)	0.92 (0.89, 0.95)
≥ 3 mo	4435	78.0 (76.3, 79.7)	1 (Ref)	76.2 (74.3, 77.9)	1 (Ref)
Breastfeeding at 3 mo, n = 10 031^d					
Leave length					
< 3 mo	6410	62.2 (60.5, 63.8)	0.87 (0.84, 0.90)	63.2 (61.5, 64.7)	0.90 (0.87, 0.94)
≥ 3 mo	3621	71.6 (69.5, 73.5)	1 (Ref)	69.8 (67.6, 71.9)	1 (Ref)

Note. APR = adjusted prevalence ratio; CI = confidence interval; PR = prevalence ratio. This analysis excluded women who did not plan to return to the same job they had during pregnancy.

^aUnweighted sample size.

^bWeighted % (95% CI).

^cAdjusted for type of leave, maternal race and Hispanic origin, age, education, marital status, household income by federal poverty level, parity, infant gestational age, timing of survey completion and Pregnancy Risk Assessment Monitoring System (PRAMS) site. Type of leave and household income by federal poverty level were obtained from PRAMS survey data. All other covariates were obtained from birth certificate data available in the PRAMS data set.

^dSample was restricted to those who had completed their PRAMS survey at or after 3 mo after giving birth.

breastfeeding at 2 and 3 months, there was only a significant interaction between leave length and maternal race and Hispanic origin (Table C [available as a supplement to the online version of this article at <http://www.ajph.org>]). Breastfeeding at 2 months was lower among women who reported less than 3 months compared with 3 or more months of leave for women who were non-Hispanic Black (64.8% vs 78.5%; APR = 0.83; 95% CI = 0.77, 0.89), non-Hispanic other (72.4% vs 82.3%; APR = 0.88; 95% CI = 0.81, 0.95), and non-Hispanic White (69.7% vs 74.6%; APR = 0.93; 95% CI = 0.90, 0.97), respectively. Breastfeeding at 3 months was lower among women who reported

taking or planning to take less than 3 months compared with 3 or more months of leave for women who were non-Hispanic Black (55.9% vs 73.5%; APR = 0.76; 95% CI = 0.69, 0.83), non-Hispanic other (63.8% vs 73.5%; APR = 0.87; 95% CI = 0.78, 0.96), and non-Hispanic White (63.6% vs 68.8%; APR = 0.92; 95% CI = 0.88, 0.97), respectively. No differences in prevalence of breastfeeding at 2 and 3 months by leave length were observed among Hispanic women (76.4% vs 76.8% at 2 months; 68.6% vs 66.9% at 3 months).

No significant interactions between leave length and type of leave were observed for breastfeeding outcomes (Table 3). Shorter leave length was

associated with lower rates of breastfeeding at 2 and 3 months, independent of whether any leave was paid.

DISCUSSION

Despite efforts to increase breastfeeding support in the workplace,¹¹ differences in breastfeeding duration were evident by length of leave. In this analysis of PRAMS data, we found that approximately two thirds of women took or planned to take less than 3 months of leave after delivery. Breastfeeding initiation was high (> 90%); however, fewer women continued to breastfeed at 1, 2, and 3 months, which is consistent with national estimates.⁵

TABLE 3— Prevalence and Prevalence Ratio of Breastfeeding Initiation and Breastfeeding at 1, 2, and 3 Months by Leave Length Stratified by Type of Leave Among Women With a Recent Live Birth Who Were Employed During and After Pregnancy: Pregnancy Risk Assessment Monitoring System, 10 US sites, 2016–2018

	Total No. ^a	No Paid Leave		Any Paid Leave		P ^e
		% (95% CI) ^{b,c}	APR (95% CI) ^{c,d}	% (95% CI) ^{b,c}	APR (95% CI) ^{c,d}	
Initiated breastfeeding, n = 12 301						
Leave length						.3
< 3 mo	7866	91.2 (90.0, 92.3)	0.98 (0.96, 1.00)	89.4 (87.9, 90.7)	0.96 (0.94, 0.98)	
≥ 3 mo	4435	93.2 (91.4, 94.6)	1 (Ref)	93.2 (91.4, 94.6)	1 (Ref)	
Breastfeeding at 1 mo, n = 12 301						
Leave length						.11
< 3 mo	7866	80.9 (79.1, 82.5)	0.96 (0.93, 1.00)	78.5 (76.7, 80.3)	0.93 (0.90, 0.96)	
≥ 3 mo	4435	83.9 (81.4, 86.1)	1 (Ref)	84.9 (82.6, 86.9)	1 (Ref)	
Breastfeeding at 2 mo, n = 12 301						
Leave length						.6
< 3 mo	7866	71.5 (69.5, 73.3)	0.93 (0.89, 0.97)	68.8 (66.8, 70.7)	0.91 (0.87, 0.95)	
≥ 3 mo	4435	76.8 (74.0, 79.3)	1 (Ref)	75.6 (73.0, 77.9)	1 (Ref)	
Breastfeeding at 3 mo, n = 10 031^f						
Leave length						.91
< 3 mo	6410	64.5 (62.2, 66.7)	0.91 (0.86, 0.96)	62.0 (59.7, 64.2)	0.90 (0.86, 0.95)	
≥ 3 mo	3621	71.2 (67.9, 74.2)	1 (Ref)	68.6 (65.7, 71.4)	1 (Ref)	

Note. APR = adjusted prevalence ratio; CI = confidence interval; PR = prevalence ratio. This analysis excluded women who did not plan to return to the same job they had during pregnancy.

^aUnweighted sample size.

^bWeighted % (95% CI).

^cAdjusted for type of leave, length of leave, maternal race and Hispanic origin, age, education, marital status, household income by federal poverty level, parity, infant gestational age, timing of survey completion and their Pregnancy Risk Assessment Monitoring System site (PRAMS). Type of leave and household income by federal poverty level were obtained from PRAMS survey data. All other covariates were obtained from birth certificate data available in the PRAMS data set.

^dWe constructed separate multivariable survey-weighted logistic regression models to examine the association with each breastfeeding outcome by type of leave between women who reported < 3 mo of leave and those who reported ≥ 3 mo of leave. The reference group for each model was ≥ 3 mo of leave.

^eP value based on F-test for 2-way interaction between leave length and type of leave.

^fSample was restricted to those who had completed their PRAMS survey at or after 3 mo after giving birth.

Women who reported taking or planning to take less than 3 months of leave were less likely to initiate and continue breastfeeding than were women with 3 or more months of leave, with estimates suggesting that this difference increased for each additional month of breastfeeding duration measured. The absolute differences in breastfeeding duration were modest (ranging from 4.8% at 1 month to 6.6% at 3 months). By 3 months, fewer than two thirds of women who had less than 3 months of

leave reported any breastfeeding. Although this finding suggests a low likelihood that women in our sample would meet the American Academy of Pediatrics breastfeeding recommendation to exclusively breastfeed to about 6 months and continue breastfeeding until 1 year or more,¹ we were unable to measure breastfeeding exclusivity or breastfeeding duration beyond 3 months.

Previous studies have yielded mixed results on the relationship between

paid leave and breastfeeding duration, with studies finding a positive or null effect on breastfeeding duration.^{15,19} We found no significant interaction between leave length and whether leave was paid for breastfeeding outcomes. Women with shorter leave length, independent of whether it was paid or unpaid, were less likely than were those with longer leave to continue breastfeeding at 2 or 3 months. However, a higher proportion of women with any paid leave reported

taking or planning to take 3 or more months of leave compared with those with no paid leave. These findings suggest that any amount of paid leave might indirectly affect breastfeeding rates by influencing the total length of leave women take. However, we were unable to examine the proportion of usual pay received while on leave, which might also influence decisions on leave length. Previous research has shown that women with paid leave are more likely to take longer leave.⁸ Some evidence suggests state-based paid leave policies might be a mechanism for enabling women who might not otherwise be able to afford to take leave to be able to take longer postpartum leave.²⁰ Of note, women in NYC, where a statewide paid leave policy was implemented in 2018,¹² reported the highest prevalence of 3 or more months of leave among PRAMS sites.

In our sample, the sociodemographic differences (e.g., age, race and Hispanic origin, education) related to the length of leave taken are similar to differences in census data findings on women who received any paid leave and longer periods of leave.⁸ Previous studies have also found differences in breastfeeding outcomes by many sociodemographic characteristics, including race and Hispanic origin.^{27,28} The significant interaction between leave length and race and Hispanic origin that we found suggests that longer leave minimizes differences in breastfeeding prevalence by race. This finding suggests that access to longer leave may be a strategy to reduce racial/ethnic disparities in breastfeeding rates. Overall, longer leave length was associated with improved breastfeeding rates among all racial/ethnic groups, except for women who were Hispanic. It is unclear

why this relationship was null; however, previous research has demonstrated that Hispanic women have higher breastfeeding rates, independent of other factors typically associated with breastfeeding rates.²⁹ In addition, it is possible that our analysis might not have been powered to detect differences among Hispanic women.

The proportion of women who take any maternity leave has remained stagnant since 1994,³⁰ despite efforts to expand access through the FMLA and state-level leave programs.^{12,20,30} Although some recent evidence suggests that uptake of the Affordable Care Act provision to cover access to lactation services and breast pumps is associated with increased breastfeeding duration,³¹ breastfeeding rates remain suboptimal.^{1,5} Despite efforts to increase support of breastfeeding in the workplace, we found that leave length was associated with breastfeeding outcomes. This association contributes to existing evidence about the role that leave plays on the ability of women with a recent live birth to meet American Academy of Pediatrics breastfeeding recommendations.

Limitations

This analysis is subject to several limitations. We did not have data on the type, location, size of the respondents' employer, work schedule (e.g., part-time, full-time, flexible schedule), or specific type of leave (vacation time, sick time, FMLA, etc.), which might also influence breastfeeding duration. Women who return to work full-time are more likely to cease breastfeeding than are women who return to work part-time.³² The type of work schedule has also been shown to

play an important role in whether women meet their breastfeeding intentions.³³

Also, data were unavailable on workplace leave taken or plans for leave for women who returned to a different job than the one they had during pregnancy and for women who were unemployed during pregnancy and actively seeking employment. Therefore, our findings might underrepresent women who had returned or were planning to return to work. In addition, PRAMS did not have data on the proportion of usual pay women received while on leave or the proportion of leave that was paid or unpaid for respondents who reported both.

We also could not examine breastfeeding-related measures that might have confounded our findings, such as breastfeeding intentions and reasons for not starting or stopping breastfeeding. We were also unable to examine breastfeeding exclusivity and any breastfeeding beyond 3 months. PRAMS data are self-reported and subject to social desirability and recall bias. Recall bias might be unlikely, as PRAMS data are collected 2 to 6 months after giving birth and most respondents in this study sample (87%) completed and returned the PRAMS survey between 2 to 4 months after giving birth.³⁴

Finally, our findings are also limited to PRAMS sites that included work-related questions on their site-specific survey, potentially limiting the generalizability of our findings to other sites. Despite these limitations, PRAMS provides a rich source of data from women with a recent live birth, which allowed us to examine associations of leave length with breastfeeding outcomes and consider important interactions.

Public Health Implications

Among women who were employed during pregnancy and returning to work after delivery, nearly all reported taking some leave, with approximately two thirds reporting less than 3 months of leave. Women reporting less than 3 months of leave were less likely to initiate breastfeeding and continue breastfeeding at 1, 2, and 3 months than were women with 3 or more months of leave. A higher proportion of women reporting any paid leave reported taking 3 or more months of leave than women reporting no paid leave. However, the association of length of leave with breastfeeding rates, in general, was independent of whether any leave was paid. Women with less than 3 months of leave reported shorter breastfeeding duration than did women with 3 or more months of leave. *AJPH*

ABOUT THE AUTHORS

Katherine Kortsmitt, Rui Li, Shanna Cox, Carrie K. Shapiro-Mendoza, Denise V. D'Angelo, Wanda D. Barfield, Holly B. Shulman, and Lee Warner are with the Division of Reproductive Health, National Center for Chronic Disease Prevention and Health Promotion, Centers for Disease Control and Prevention, Atlanta, GA. Cria G. Perrine is with the Division of Nutrition, Physical Activity, and Obesity, National Center for Chronic Disease Prevention and Health Promotion. Craig F. Garfield is with the Departments of Pediatrics and Medical Social Sciences, Northwestern University Feinberg School of Medicine and Lurie Children's Hospital of Chicago, Chicago, IL.

CORRESPONDENCE

Correspondence should be sent to Katherine Kortsmitt, PhD, MPH, Division of Reproductive Health, National Center for Chronic Disease Prevention and Health Promotion, Centers for Disease Control and Prevention, 4770 Buford Hwy NE, Building 107, Mailstop S107-2, Atlanta, GA 30341 (e-mail: nlv2@cdc.gov). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

K. Kortsmitt conceptualized the study, conducted the data analyses, and drafted the article. R. Li, S. Cox, C. K. Shapiro-Mendoza, C. G. Perrine, D. V. D'Angelo, W. D. Barfield, H. B. Shulman, C. F. Garfield, and L. Warner provided significant input on interpretation of the data and review and editing of the article.

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CONFLICTS OF INTEREST

The authors have no conflicts of interest to disclose.

HUMAN PARTICIPANT PROTECTION

This was a secondary data analysis and was exempt from institutional review board review. The Centers for Disease Control and Prevention and each site's institutional review boards reviewed and approved the Pregnancy Risk Assessment Monitoring System study protocol.

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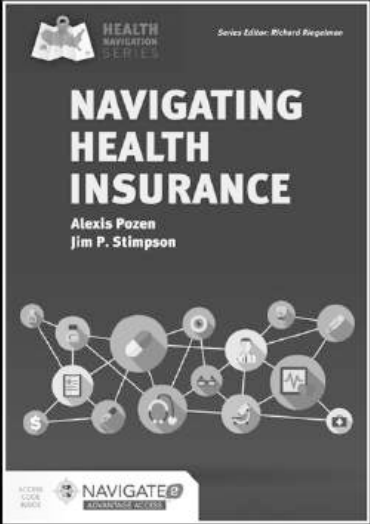
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Trends in Ketamine Use, Exposures, and Seizures in the United States up to 2019

Joseph J. Palamar, PhD, MPH, Caroline Rutherford, MS, and Katherine M. Keyes, PhD

Objectives. To determine whether there have been shifts in nonmedical ketamine use, poisonings (“exposures”), and seizures.

Methods. We used generalized additive models to detect trends in past-year use (2006–2019), exposures (1991–2019), and seizures (2000–2019) involving ketamine in the United States.

Results. There was a quarterly increase in self-reported past-year nonmedical ketamine use in 2006 to 2014 ($B = 0.21$; $P = .030$) and an increase in 2015 to 2019 ($B = 0.29$; $P = .036$), reaching a peak of 0.9% in late 2019. The rate of exposures increased from 1991 through 2019 ($B = 0.87$; $P = .006$), and there was an increase to 1.1 exposures per 1 000 000 population in 2014, with rates remaining stable through 2019. The rate of ketamine seizures increased from 2000 through 2019 ($B = 2.27$; $P < .001$), with seizures reaching a peak in 2019 at 3.2 per 1000 seizures.

Conclusions. Indicators suggest that ketamine use and availability has increased, including before increased medical indications, but nonmedical use is still currently uncommon despite increased acceptance and media coverage. (*Am J Public Health.* 2021;111(11):2046–2049. <https://doi.org/10.2105/AJPH.2021.306486>)

Ketamine is a dissociative anesthetic that has been prevalent as a recreational drug in nightclubs for decades.^{1,2} However, ketamine has been used as an anesthetic in both animals and humans for a half century.³ Evidence indicates that ketamine produces analgesia without substantial respiratory depression in both children and adults, and thus its use is particularly common in prehospital settings (e.g., among emergency medical technicians) and in field conditions (e.g., in low-income countries).³ Ketamine has not been widely used in psychiatry until recently. Randomized double-blind controlled-placebo trials have recently determined that intranasal nonanesthetic doses of esketamine, an enantiomer of ketamine, can have an ultrarapid antidepressant effect among

those with major depression.⁴ In response to studies demonstrating the drug’s efficacy in treating treatment-resistant depression, on March 5, 2019, the US Food and Drug Administration approved the use of esketamine nasal spray to treat this condition.

Findings on the drug’s efficacy and Food and Drug Administration approval led to increased availability of ketamine in psychiatric treatment settings. Such coverage, however, also led to extensive media coverage of the drug. Since ketamine’s approval for use in psychiatric settings, the drug has been covered in many major US news sources, including *USA Today* and the *New York Times*,^{5,6} and many articles refer to ketamine as a “club drug” or refer to it by its street name: “Special K.” Given that media coverage can affect drug use patterns in the

general population,⁷ it is unknown whether such coverage about the benefits of use of a club drug may have created mixed messages regarding the drug’s safety or acceptability to use recreationally. Estimated past-year use of ketamine—which is assumed to be nonmedical use—was found to have significantly increased among nightclub and dance festival attendees in New York City between 2016 and 2019 (from 5.9% to 15.3%),² but trends in use in the general population have remained largely unknown. We examined trends in ketamine use and availability using multiple forms of national indicator data.

METHODS

We analyzed data from 3 national data sources. First, we estimated trends in

quarterly prevalence of ketamine use based on past-year report among participants aged 12 to 34 years in the 2006 to 2019 National Survey on Drug Use and Health (NSDUH), a nationally representative survey of noninstitutionalized individuals in the United States.⁸ Ketamine use was asked about in a section querying hallucinogen or psychedelic use in which participants were asked about use of “ketamine, also called ‘Special K’ or ‘Super K,’” so we assumed reported use to be mainly nonmedical use.

We examined quarterly trends in past-year use separately (using weighted data) before and after 2015 because of a change in survey design.⁸ Then, we estimated trends in poisonings (“exposures”) reported to poison control centers (PCCs) involving ketamine and its analogs from 1991 to 2019. There are 55 PCCs that cover the United States and its territories, and, since 2003, data have been stored in the National Poison Data System, which replaced the previous Toxic Exposure Surveillance System.⁹ We converted counts to rates per 1 000 000 persons based on the US Census. Finally, we estimated trends in seizures tracked by the Drug Enforcement Administration’s National Forensic Laboratory Information System (NFLIS) from 2000 to 2019 based on counts converted into rates per 1000 annual total drug seizures.

The NFLIS systematically collects seizure results from federal, state, and local forensic laboratories throughout all 50 states.¹⁰ We chose years for inclusion for all data sources based on data availability. We analyzed case-level data for NSDUH and extracted count data from PCCs and the NFLIS from annual reports.^{9,10} We used generalized additive models with cubic basis functions to fit regression splines with

automated selection of knots to visually capture nonlinear trends. We plotted trends with 95% confidence intervals for model predictions. We then modeled trends with orthogonal polynomial terms. Data and code can be found at <https://github.com/caroruth/ketaminetrends>.

RESULTS

There was a linear quarterly increase in self-reported past-year ketamine use (via the NSDUH) in 2006 to 2014 ($B = 0.21$; $SE = 0.01$; $P = .030$) with a nadir in late 2008 of 0.1% and then a cubic increase in 2015 to 2019 ($B = 0.29$; $SE = 0.13$; $P = .036$), reaching a peak of 0.9% in late 2019 (Figure 1). The rate of exposures reported to PCCs increased in a cubic manner from 1991 through 2019 ($B = 0.87$; $SE = 0.29$; $P = .006$; Figure A, available as a supplement to the online version of this article at <https://www.ajph.org>). Reported exposures began in 1991 at 0.1 exposures per 1 000 000 population and increased to a peak of 1.4 exposures per 1 000 000 in 2000. Rates then decreased to 0.4 exposures per 1 000 000 in 2008 and increased to 1.1 exposures per 1 000 000 in 2014 and remained somewhat stable through 2019. The rate of seizures increased in a quadratic manner from 2000 through 2019 ($B = 2.27$; $SE = 0.33$; $P < .001$), with seizures reaching a nadir in 2004 to 2005 at 0.27 per 1000 and increasing from 2011 through a peak in 2019 at 3.2 per 1000 (Figure B, available as a supplement to the online version of this article at <https://www.ajph.org>).

DISCUSSION

Our results suggest increased nonmedical use and availability of ketamine in recent years—particularly in the past

year. However, despite a recent increase in prevalence of nonmedical ketamine use, prevalence remained relatively rare, below 1%. Exposures as reported to PCCs were highest in 2000 to 2001, which was also the peak in use of other club drugs such as MDMA (3,4-methylenedioxymethamphetamine), also known as ecstasy,¹¹ possibly suggesting ketamine’s close link to recreational use in nightlife scenes. After a decline in exposures, there was an increase again through 2014, but exposures have remained relatively consistent thereafter. Rates of ketamine seizures, however, have increased exponentially since 2012, suggesting increasing availability, although seizures are still uncommon relative to other drugs. Overall, results suggest that use and availability are indeed increasing, but ketamine remains an uncommon recreational drug.

Detected increases in ketamine use corroborate recent detected increases in recreational use among nightclub and dance festival attendees in New York City,² so use appears to be increasing both in this high-risk population and in the general population. However, to better inform prevention and harm reduction efforts, future research is needed to determine use and exposure trends according to demographic and other drug use characteristics to obtain a clearer picture regarding which subpopulations are increasing use. Further, studies are needed to directly determine whether findings on the efficacy of ketamine and associated media coverage are directly linked to increasing use.

Limitations

This study is not without limitations. Data on ketamine use and exposure

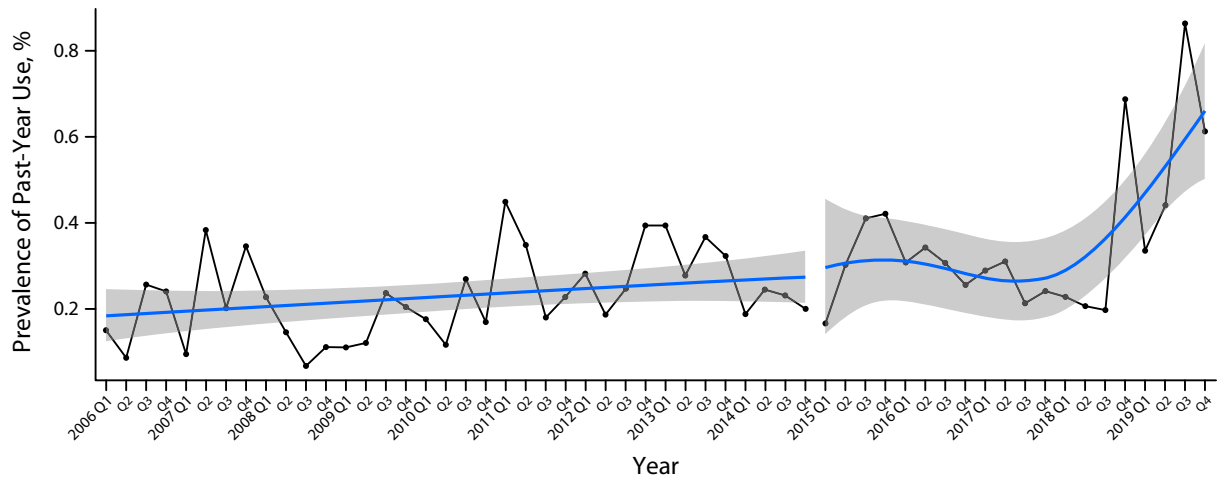


FIGURE 1— Trends in Self-Reported Past-Year Ketamine Use, Aged 12–34 Years: United States, 2006–2019

are likely underreported because data rely on self-report. We assumed NSDUH data on ketamine use to refer to nonmedical use, but it is possible that some use was medical use. Count data from PCCs was limited, as we were not able to deduce with confidence the extent of “abuse” or intentional misuse among cases. Finally, people who use synthetic drugs such as ecstasy can be unknowingly exposed to ketamine as an adulterant, and this also leads to underreporting of use.¹²

Public Health Implications

National indicator data suggest that nonmedical ketamine use and availability are increasing in the United States, but use is still uncommon—even despite increased acceptance and media coverage. Ketamine is used medically throughout much of the world, and access to appropriate medical use is warranted; however, it is important for surveillance efforts focusing on nonmedical use to also continue to inform prevention and harm reduction. **AJPH**

ABOUT THE AUTHORS

Joseph J. Palamar is with the Department of Population Health, New York University Grossman School of Medicine, New York, NY. Caroline Rutherford and Katherine M. Keyes are with the Department of Epidemiology, Mailman School of Public Health, Columbia University, New York, NY.

CORRESPONDENCE

Correspondence should be sent to Joseph J. Palamar, New York University Grossman School of Medicine, Department of Population Health, New York, NY 10016 (e-mail: joseph.palamar@nyulangone.org). Reprints can be ordered at <http://www.ajph.org> by clicking the “Reprints” link.

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CONTRIBUTORS

J.J. Palamar conceptualized and designed the study. C. Rutherford conducted the statistical analyses under the mentorship of K. M. Keyes. All authors drafted the initial article, interpreted results, and critically reviewed and revised the article. All authors are responsible for this reported research and approved the final article as submitted.

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CONFLICTS OF INTEREST

J.J. Palamar has consulted for the biopharmaceutical company Alkermes plc. The authors have no other potential conflicts of interest to declare.

HUMAN PARTICIPANT PROTECTION

This study was exempt from institutional review board review because it was a secondary data analysis and therefore did not include human participants.

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Racial Disparities of E-Cigarette Use Among US Youths: 2014–2019

Hongying Dai, PhD, Athena K. Ramos, PhD, Babalola Faseru, MD, MPH, Jennie L. Hill, PhD, and Steven Y. Sussman, PhD

Objectives. To evaluate disparities in youth e-cigarette use patterns and flavor use by race/ethnicity over time.

Methods. We used data from the US 2014–2019 National Youth Tobacco Survey (NYTS) to examine trends in dual use (co-use of e-cigarettes and cigarettes or other tobacco products), occasional (≤ 5 days) versus frequent use (≥ 20 days) in the past 30 days, and flavor use among current (past-30-day) e-cigarette users ($n = 13\,178$) across racial/ethnic groups (non-Hispanic Whites, non-Hispanic Blacks, Hispanics/Latinos, and non-Hispanic others).

Results. Among current e-cigarette users, dual use and occasional use decreased significantly from 2014 to 2019 across racial and ethnic groups except for non-Hispanic Blacks; frequent use and flavored e-cigarette use increased among non-Hispanic Whites, Hispanics/Latinos, and non-Hispanic others but not among non-Hispanic Blacks. In 2019, non-Hispanic Black e-cigarette users were more likely to report dual use (adjusted odds ratio [AOR] = 2.2; 95% confidence interval [CI] = 1.5, 3.2; $P < .001$) and occasional use of e-cigarettes (AOR = 3.7; 95% CI = 2.3, 5.9; $P < .001$) but less likely to report frequent use (AOR = 0.2; 95% CI = 0.1, 0.4; $P < .001$) and flavored e-cigarette use (AOR = 0.4; 95% CI = 0.3, 0.5; $P < .001$) than their White peers.

Conclusions. Youth e-cigarette use patterns differed considerably across racial/ethnic groups, and tailored strategies to address disparities in e-cigarette use are needed. (*Am J Public Health.* 2021;111(11):2050–2058. <https://doi.org/10.2105/AJPH.2021.306448>)

While the cigarette smoking rate among youths has been declining over the past several decades,^{1,2} the prevalence of current (past-30-day) e-cigarette use (or vaping) among adolescents increased dramatically during 2017 to 2019.^{2,3} Although e-cigarettes deliver a substantially lower level of toxins than do combustible cigarettes,⁴ e-cigarette aerosol is not harmless, as studies have identified harmful and potentially harmful constituents in e-cigarettes.⁵ Vaping at a young age could cause nicotine addiction, harm brain development, and increase risks of respiratory and cardiovascular diseases.^{5–7}

Race/ethnicity may differentiate youth tobacco use, and non-Hispanic Black and Hispanic/Latino tobacco users experience significant health disparities in tobacco-related diseases, including cancer, stroke, and heart disease.^{8,9} Studies have documented that non-Hispanic Black adolescents report a significantly lower prevalence of cigarette smoking but a higher prevalence of cigar use than their Hispanic/Latino and White peers.¹⁰ Hispanic/Latino youths reported a higher prevalence of hookah use.¹⁰ Recent studies also indicate that e-cigarettes are not uniformly used across racial and ethnic groups,⁵ and non-Hispanic White and Hispanic/

Latino youths are more likely to use e-cigarettes than are non-Hispanic Black youths.^{10,11} Longitudinal studies also show distinct transition patterns by racial and ethnic group,^{12,13} and White and Hispanic adolescent e-cigarette users are more likely to transition to cigarette smoking than their Black counterparts.¹³ Another national study found that non-Hispanic Black students and Hispanic students significantly initiate e-cigarettes at an earlier age than White peers.¹⁴ Alarmingly, early initiation (age 13 years and younger vs older than 13 years) of e-cigarettes could increase the risk of nicotine dependence and sustained

e-cigarette use.¹⁴ A socioecological model posits that multifaceted factors at the individual, interpersonal, community, and policy levels could lead to distinct patterns of exposure to secondhand tobacco, tobacco initiation, use patterns, and cessation behaviors by race and ethnicity.¹⁵ These existential disparities in e-cigarette and tobacco use may be attributable to social determinants related to race/ethnicity (e.g., education, income, geography), neuropsychological factors, the long history of the tobacco industry's aggressive marketing toward racial and ethnic minority communities, and biological aspects.^{5,8,9}

A growing body of literature has extended current youth e-cigarette research to better understand the frequency of e-cigarette use, co-use of e-cigarettes and other tobacco products, and flavored e-cigarette use. For instance, Glasser et al.¹⁶ reported that about half of current e-cigarette users vaped occasionally (≤ 5 days in the past 30 days) and roughly a quarter vaped frequently (≥ 20 days in the past 30 days) in 2018.¹⁶ E-cigarette use is strongly associated with cigarette smoking and other tobacco use; a majority of young current e-cigarette users have been found to report concurrently using 1 or more other tobacco products.^{5,17} A previous study¹⁷ identified that, across racial and ethnic groups, most current tobacco users were dual or poly tobacco users in 2014. But it is unclear whether poly tobacco use behaviors have changed since then.

Meanwhile, flavored e-cigarette products are widely available in the United States, and flavors have become one of the leading reasons for current e-cigarette use among youths.¹⁸ Flavored e-cigarette use has been

increasing among youths since 2015, with 65.2% of current e-cigarette users reporting use of flavored products in the past 30 days in 2018.^{19,20} However, little is known about differential patterns of e-cigarette use and flavor use across racial/ethnic groups over time. Understanding how adolescents use e-cigarettes (e.g., use patterns, flavors) is critical to inform regulatory actions and develop effective intervention strategies to prevent and reduce youth vaping behaviors, especially among vulnerable subpopulations such as racial and ethnic minority adolescents.

To address these gaps in knowledge, we analyzed data from the 2014–2019 National Youth Tobacco Survey (NYTS) by examining the changes in temporal trends of e-cigarette use patterns and flavor use among current youth e-cigarette users across racial and ethnic groups. We further tested disparities in e-cigarette use patterns and flavor use across racial/ethnic groups.

METHODS

The NYTS is a nationally representative, cross-sectional, and school-based annual survey of middle- and high-school students between the ages of 9 and 19 years in the United States. The survey was conducted using a probabilistic sampling procedure without replacement at 3 stages: (1) primary sampling units such as a county, a group of small counties, or part of a very large county; (2) secondary sampling units including schools within each selected primary sampling unit; and (3) students within each selected school. The 2014–2019 NYTS data included 117 472 respondents, with the annual survey sample size ranging from 17 711 in 2015 to 22 007 in 2014. The median response rate for participating

schools and students ranged from 63.4% to 73.3% during the study period. A detailed description of the NYTS survey can be found on the NYTS Web site.²¹

Measures

Race/ethnicity. Participants were classified into 4 groups: non-Hispanic White, non-Hispanic Black, Hispanic/Latino, or non-Hispanic other.

E-cigarette use. Current e-cigarette users were determined by those who reported using e-cigarettes 1 or more days in the past 30 days. We further categorized current e-cigarette users as occasional users (≤ 5 days) and frequent users (≥ 20 days) based on the frequency of e-cigarette use in the past 30 days.

Current cigarette smokers were defined as those who reported smoking 1 or more days in the past 30 days, and current other tobacco users were defined as those who reported using 1 or more other tobacco products on 1 or more days of the past 30 days. Other tobacco products included cigars (cigars, little cigars, and cigarillos), smokeless tobacco (chewing tobacco, snuff, dip, snus, and dissolvable tobacco), hookahs, pipe tobacco, and bidis.¹⁰ Those who reported current co-use of e-cigarettes and cigarettes or other tobacco products were defined as dual users. Current e-cigarette users who reported using e-cigarettes that tasted like menthol (mint), alcohol (wine, cognac), candy, fruit, chocolate, or any other flavors were classified as flavored e-cigarette users. The NYTS did not have separate questions for each flavor, and a single composite flavor question was utilized. Those who reported not using flavored

e-cigarettes were likely to be users of tobacco-flavored or flavorless e-cigarettes or those who did not recall the flavor in e-cigarettes.

Covariates. Demographic variables include gender (male or female), age (continuous variable), school level (middle or high school), and tobacco use by other household members (“none,” “other tobacco product use” [i.e., non-e-cigarette use], and “e-cigarettes”).

Statistical Analyses

We weighted data to provide national estimates by accounting for the complex survey design and nonresponse. First, weighted percentages and 95% confidence intervals (CIs) of sample characteristics (e.g., demographic characteristics, smoking prevalence) among current e-cigarette users were reported from 2014 to 2019.

Second, temporal trends of e-cigarette use patterns (dual use, occasional use, and frequent use) and flavored e-cigarette use from 2014 to 2019 were reported, overall and by racial/ethnic groups. We examined the interaction of year \times race/ethnicity in the multivariable model adjusted for covariates, current cigarette smoking, and other tobacco use status. We also reported linear and quadratic trends based on multivariable logistic regression analyses, where survey years served as a continuous variable. We performed stratified analyses by racial/ethnic group. The sample sizes for other races, such as Asians, American Indians/Alaska Natives, Native Hawaiians, and other Pacific Islanders, were too small to model the temporal trends. Therefore, we combined these minority subpopulations into 1 category (i.e., non-Hispanic others).

Third, we performed separate multivariable logistic regression models to examine the association between race/ethnicity and e-cigarette use patterns and flavor use among current e-cigarette users by using the 2019 NYTS data. The temporal trends in associations between race/ethnicity and e-cigarette use patterns from 2014 to 2019 are reported in Figure A (available as a supplement to the online version of this article at <http://www.ajph.org>). Adjusted odds ratios (AORs) are reported in the multivariable analysis. We analyzed data by using SAS version 9.4 (SAS Institute, Cary, NC) survey procedures, and we considered a *P* level of less than .05 to be statistically significant.

RESULTS

The analytical sample included 13 178 current e-cigarette users from the combined 2014–2019 NYTS (female, 44.6%; high school, 79.8%; non-Hispanic White, 65.0%; non-Hispanic Black, 7.6%; Hispanic/Latino, 23.6%; current cigarette smokers, 28.6%; current other tobacco use, 43.0%). Table A (available as a supplement to the online version of this article at <http://www.ajph.org>) presents the sample characteristics of current e-cigarette users from 2014 to 2019. Overall, the distribution of e-cigarette users by age and grade was relatively stable, while more female students were current e-cigarette users in 2019 (47.8%) than in 2014 (44.1%). Among current e-cigarette users, the prevalence of current cigarette smoking dropped dramatically from 37.6% in 2014 to 17.9% in 2019. We also observed a decreasing trend in the prevalence of current other tobacco use (56.5% in 2014 to 30.8% in 2019). Table B reports the unweighted sample

sizes for each outcome variable by racial/ethnic group.

Figure 1 presents the temporal trends of dual use among current e-cigarette users from 2014 to 2019. The overall dual use decreased significantly from 65.2% in 2014 to 36.0% in 2019 (difference = -29.2% ; 95% CI = -34.4% , -24.0%) and fit a significant quadratic trend over 2014 to 2019 ($P = .009$). Changes in dual use among current e-cigarette users from 2014 to 2019 differed significantly by racial and ethnic group (year \times race/ethnicity; $P < .001$). The decreases were significant for non-Hispanic Whites (AOR = 0.76; 95% CI = 0.72, 0.80; $P < .001$) and non-Hispanic others (AOR = 0.83; 95% CI = 0.71, 1.00; $P = .046$). The changes fit a quadratic trend for Hispanics/Latinos ($P < .001$). For non-Hispanic Blacks, there was no significant change in dual use from 48.3% in 2014 to 49.9% in 2019 (AOR = 0.97; 95% CI = 0.88, 1.06; $P = .46$).

Among current e-cigarette users, there was a significant decrease in occasional use (Figure 2a; 63.5% in 2014 vs 50.8% in 2019; quadratic trend, $P < .001$) and a significant increase of frequent use (Figure 2b; 14.8% in 2014 vs 30.4% in 2019; quadratic trend, $P < .001$). Changes in use patterns among current e-cigarette users differed by racial and ethnic group (year \times race/ethnicity; $P < .001$). The drops in occasional use were significant across all racial/ethnic groups except for non-Hispanic Blacks, for whom the proportion was stable at 72.2% in 2014 versus 71.2% in 2019 (AOR = 1.01; 95% CI = 0.90, 1.13; $P = .87$). Similarly, the increases in frequent use were significant across all racial/ethnic groups except for non-Hispanic Blacks with a stable trend at 11.8% in 2014 versus

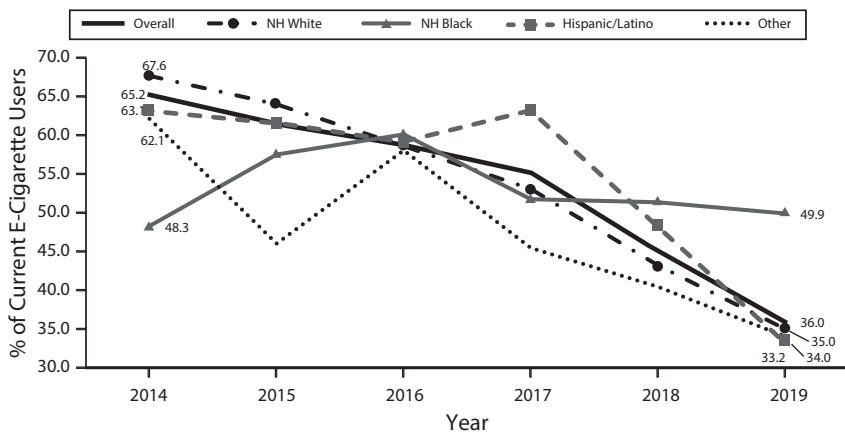


FIGURE 1— Trends in Dual Use Among Current E-Cigarette Users: United States, 2014-2019

Note. NH = non-Hispanic. The y-axis shows the observed % of dual use (co-use of e-cigarettes and cigarettes or other tobacco products) among current e-cigarette users. Multivariable regression models adjusted for age, gender, and tobacco use by a household member. The interaction for year \times race/ethnicity was $P < .001$. Overall: quadratic trend, $P = .009$. NH White: linear trend, adjusted odds ratio (AOR) = 0.76; 95% confidence interval (CI) = 0.72, 0.80; $P < .001$; quadratic trend, $P = .24$. NH Black: linear trend, AOR = 0.96; 95% CI = 0.87, 1.06; $P = .40$; quadratic trend, $P = .16$. Hispanic/Latino: quadratic trend, $P < .001$. NH other: linear trend, AOR = 0.84; 95% CI = 0.71, 1.00; $P = .046$; quadratic trend, $P = .67$.

13.5% in 2019 (AOR = 1.10; 95% CI = 0.95, 1.27; $P = .19$).

Changes in flavor use among current e-cigarette users fit a quadratic trend from 2014 to 2019 ($P < .001$) with a significant decrease from 2014 to 2015 and then a significant increase from 2015 to 2019 (Figure 3). The interaction term between the year and race/ethnicity was not significant ($P = .23$).

Table 1 presents univariate and multivariable analyses of e-cigarette use patterns and race/ethnicity among current e-cigarette users in 2019 ($n = 3628$). Non-Hispanic Black e-cigarette users had higher odds of reporting dual use compared with their White peers (AOR = 2.2; 95% CI = 1.5, 3.2; $P < .001$). Non-Hispanic White e-cigarette users were less likely than their racial/ethnic minority peers to report occasional use but more likely to report frequent use. For instance, 44.7% (95% CI = 41.6%, 47.8%) of non-Hispanic White e-cigarette users reported occasional use in 2019, compared with 71.2%

(95% CI = 64.0%, 78.4%) of non-Hispanic Blacks (AOR = 3.7; 95% CI = 2.3, 5.9; $P < .001$) and 59.0% (95% CI = 54.9%, 63.1%) of Hispanics/Latinos (AOR = 1.6; 95% CI = 1.3, 2.0; $P < .001$). Conversely, 36.3% (95% CI = 32.9%, 39.7%) of non-Hispanic White e-cigarette users reported frequent use in 2019, in comparison with 13.5% (95% CI = 8.4%, 18.6%) of non-Hispanic Blacks (AOR = 0.2; 95% CI = 0.1, 0.4; $P < .001$) and 21.2% (95% CI = 17.3%, 25.1%) of Hispanics/Latinos (AOR = 0.5; 95% CI = 0.4, 0.7; $P < .001$).

Flavored e-cigarette use differed significantly by racial and ethnic group, with non-Hispanic White e-cigarette users reporting the highest proportion (75.0%; 95% CI = 72.5%, 77.6%) and non-Hispanic Blacks reporting the lowest proportion (43.9%; 95% CI = 37.1%, 50.7%; AOR = 0.4; 95% CI = 0.3, 0.5; $P < .001$).

Temporal trends of associations between e-cigarette use behaviors and race/ethnicity are presented in

Figure A. The sensitivity analysis shows wide gaps in dual use, occasional use, and frequent use between non-Hispanic Blacks and non-Hispanic Whites.

DISCUSSION

Although enormous progress has been made in reducing tobacco use in the United States, this progress has not been equally distributed across the population, with a large disparity in tobacco use persisting across groups defined by race/ethnicity, education level, income level, region, and other factors.¹⁵ The tobacco use landscape has substantially changed in recent years with more adolescents using e-cigarettes, and the prevalence of youth e-cigarette use has surpassed the use of cigarettes since 2015.²² However, there is a dearth of research in assessing disparities of e-cigarette use with little work comparing trajectory and etiology in e-cigarette use patterns by racial and ethnic subpopulation. Surveillance of current e-cigarette use is important for public health, but attention to past-30-day e-cigarette use prevalence may obscure important use behavioral differences such as occasional use versus frequent use and co-use of e-cigarettes and other tobacco products across racial and ethnic groups.¹⁶

This study identified sharply different trends in the dual use between non-Hispanic Blacks and other racial/ethnic groups from 2014 to 2019. For instance, the proportion of current e-cigarette users reporting dual use dropped by nearly half from 65.2% in 2014 to 36.0% in 2019 for non-Hispanic Whites, but the proportion was stable for non-Hispanic Blacks (48.3% vs 49.9%). The multivariable

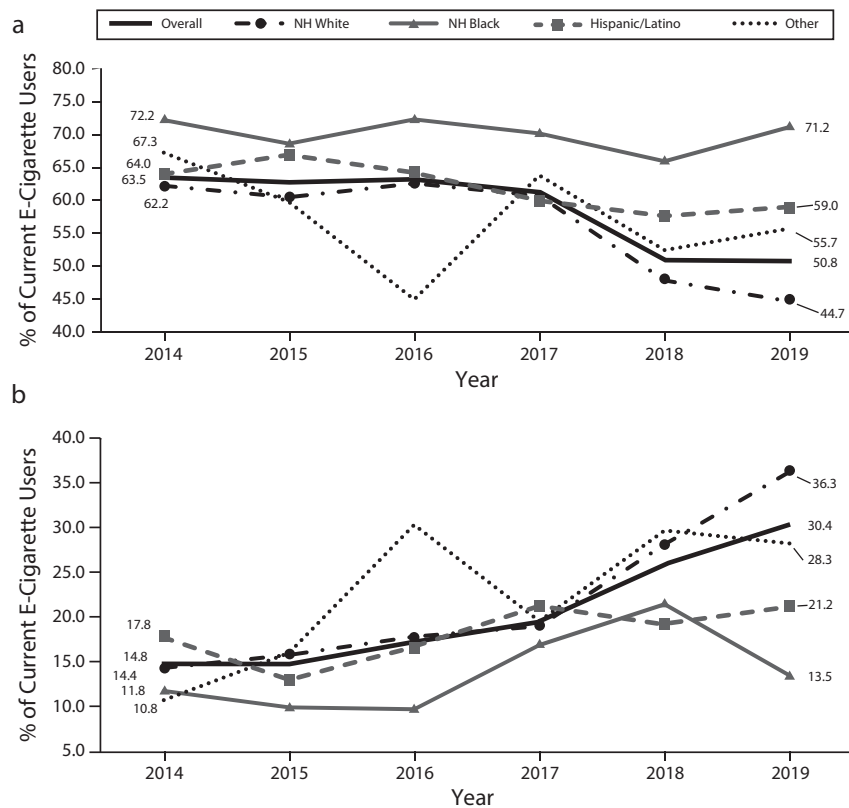


FIGURE 2— Trends in E-Cigarette Use Among Current E-Cigarette Users, by (a) Occasional Use and (b) Frequent Use: United States, 2014–2019

Note. NH = non-Hispanic. The y-axis shows the observed % of (a) occasional (≤ 5 days in the past 30 days) or (b) frequent (≥ 20 days in the past 30 days) e-cigarette use among current e-cigarette users. Multivariable regression models adjusted for age, gender, current cigarette smoking, current other tobacco use, and tobacco use by a household member. The interaction for year \times race/ethnicity was $P < .001$. Overall: quadratic trend, $P < .001$. NH White: quadratic trend, $P < .001$. NH Black: linear trend, adjusted odds ratio (AOR) = 1.01; 95% confidence interval (CI) = 0.91, 1.13; $P = .87$; quadratic trend, $P = .92$. Hispanic/Latino: quadratic trend, $P = .02$; NH others: linear trend, AOR = 0.87; 95% CI = 0.77, 0.98; $P = .02$; quadratic trend, $P = .38$. Overall: quadratic trend, $P < .001$. NH Whites: quadratic trend, $P < .001$. NH Black: linear trend, AOR = 1.10; 95% CI = 0.95, 1.27; $P = .19$; quadratic trend, $P = .92$. Hispanic/Latino: quadratic trend, $P = .005$. NH other: linear trend, AOR = 1.42; 95% CI = 1.20, 1.67; $P < .001$; quadratic trend, $P = .64$.

analysis further showed a considerable disparity, with non-Hispanic Black e-cigarette users being 1.8 times more likely to report dual use than their White peers in 2019.

Sensitivity analyses also showed a growing gap between non-Hispanic Blacks and non-Hispanic Whites in dual use over time. One may speculate that the relative preference for combustible products among non-Hispanic Blacks, possibly facilitated by the tobacco industry's targeted advertising,²³ including the

promotion of menthol cigarettes and flavored little cigars, drives these results. Another factor to consider is relatively easy access to single cigarettes and cigarillos such as Phillies, Black & Mild, and Swisher Sweets sold individually in segregated Black neighborhoods.²⁴ Dual use may further increase disparities in health outcomes given that multiple tobacco product use is associated with increased symptoms of nicotine dependence and addiction in comparison with single product use.²⁵

Our study showed that e-cigarette use patterns differed considerably by race/ethnicity. From 2014 to 2019, frequent use of e-cigarettes showed a 2.5-fold increase for non-Hispanic White e-cigarette users (14.8% to 36.3%), but the proportion was relatively flat for non-Hispanic Blacks (11.8% to 13.5%; $P = .15$). Furthermore, non-Hispanic White e-cigarette users were more likely to report frequent e-cigarette use and flavored e-cigarette use than non-Hispanic Blacks and Hispanics/Latinos in 2019. Studies have shown that vaping more frequently among adolescents was associated with a higher risk of frequent and heavy smoking in the future²⁶ and that flavored e-cigarette use was associated with an increased risk of cigarette smoking and future vaping.²⁷ Even occasional e-cigarette use is associated with significantly higher risks of binge drinking, marijuana use, and other illicit drug use than nonuse.²⁸

These remarkable disparities in e-cigarette use behaviors underscore the importance of developing and implementing tailored strategies to address the e-cigarette use epidemic across race/ethnicity. Our results suggest that non-Hispanic Black adolescents are more likely to be dual users of e-cigarette and other tobacco products, which may lead to less frequent use of e-cigarettes. For non-Hispanic Black youths, tailored interventions are needed to promote prevention and cessation of tobacco use and raise the harm perception of light tobacco use, including the occasional use of e-cigarettes. On the other hand, non-Hispanic White youths seem to be moving toward more exclusive and frequent e-cigarette use. Therefore, evidence-based youth vaping cessation programming may be suitable for this

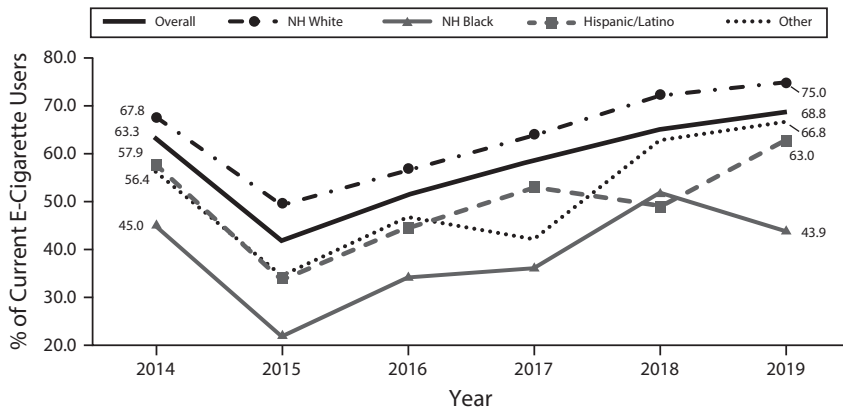


FIGURE 3— Trends in Flavored E-Cigarette Use Among Current E-Cigarette Users: United States, 2014–2019

Note. NH = non-Hispanic. The y-axis shows the observed % of flavored e-cigarette use among current e-cigarette users. Multivariable regression models adjusted for age, gender, current cigarette smoking, current other tobacco use, tobacco use by a household member, and the number of days using e-cigarettes in the past 30 days. The interaction term of year \times race/ethnicity: $P = .56$. Overall: quadratic trend, $P < .001$. NH White: quadratic trend, $P < .001$. NH Black: linear trend, adjusted odds ratio = 1.12; 95% confidence interval = 1.03, 1.22; $P = .01$; quadratic trend, $P = .11$. Hispanic/Latino: quadratic trend, $P < .001$. NH other: quadratic trend, $P = .004$.

subpopulation. Certainly, while possibly acknowledging differential preferences for various tobacco products, a uniform message might be provided—that any nicotine-containing product is dangerous for youths and may lead to cardiovascular and carcinogenic consequences later on.⁵

Enactment and implementation of tobacco control policies need to account for differential effects on racial/ethnic subpopulations to ensure that policies can uniformly prevent the initiation and reduce the prevalence of youth tobacco use, which could further lead to a reduction of health disparities in tobacco use and tobacco-related morbidity and mortality. There are large variations in tobacco-free and smoke-free public policies, tobacco taxes, tobacco retail and vape shop density, tobacco product point-of-sale advertising restrictions, and other tobacco control legislation across states and localities. In addition, the contents and coverage of e-cigarette regulations vary

by jurisdiction, which could exacerbate inequities in e-cigarette use and tobacco-related disease burden by geography, race, and ethnicity. The nationwide Tobacco 21 policy that was passed in December 2019 to raise the minimum legal age of tobacco sales to 21 years²⁹ could increase tobacco control coverage in minority populations and reduce youth access to tobacco products through commercial and social sources.

The US Food and Drug Administration issued an enforcement policy on unauthorized cartridge-based e-cigarette flavors other than tobacco and menthol, which went into effect on February 6, 2020,³⁰ and further sent warning letters to notify e-cigarette companies including Puff Bar to remove flavored disposable e-cigarette products from the market in July 2020.³¹ As menthol-flavored tobacco products are particularly appealing to African American tobacco users, continued surveillance of current e-cigarette

use by racial and ethnic groups is critically needed.

Limitations

This study had several limitations. First, the NYTS is a survey of middle- and high-school students, and our findings might not be generalizable to the broader youth population. However, 97% of adolescents aged between 10 and 17 years were enrolled in school.³²

Second, e-cigarette use behaviors were self-reported, and they are subject to recall bias, especially for younger respondents. However, the test and retest reliability of self-reported behaviors related to tobacco use among adolescents is high.³³

Third, self-reported race/ethnicity was grouped into 4 categories to ensure sufficient sample sizes in assessing temporal trends of e-cigarette use. However, a deeper understanding of differences within these groups is needed.

Fourth, these nationally representative data were of repeated cross-sections, and changes in cohort composition over time could affect results (e.g., trends between e-cigarette use and dual use). However, the sample sizes at each time point were large, and sampling strategies for data collection minimize that possibility.³⁴

Fifth, the NYTS only asked a single composite question about flavor use. It is possible that if separate flavors were recorded, the results might differ. In addition, occasional e-cigarette users may be less likely to know if the e-cigarette products they used were flavored, potentially leading to misclassified responses.

Finally, there were small changes in the wording and placement of survey questions for certain tobacco products

TABLE 1— Univariate and Multivariable Analysis of E-Cigarette Use Behaviors Among Current E-Cigarette Users: United States, 2019 National Youth Tobacco Survey

	Weighted % (95% CI)	OR (95% CI)	AOR (95% CI)
Dual use^a			
Race/ethnicity			
NH White	35.0 (30.2, 39.8)	1 (Ref)	1 (Ref)
NH Black	49.9 (41.6, 58.2)	1.8 (1.3, 2.6)	2.2 ^b (1.5, 3.2)
Hispanic/Latino	33.2 (29.4, 37.0)	0.9 (0.7, 1.2)	1.1 ^b (0.9, 1.4)
Other	34.0 (22.1, 46.0)	1.0 (0.6, 1.7)	1.0 ^b (0.6, 1.8)
Occasional e-cigarette use^c			
Race/ethnicity			
NH White	44.7 (41.6, 47.8)	1 (Ref)	1 (Ref)
NH Black	71.2 (64.0, 78.4)	3.1 (2.1, 4.5)	3.7 ^d (2.3, 5.9)
Hispanic/Latino	59.0 (54.9, 63.1)	1.8 (1.5, 2.2)	1.6 ^d (1.3, 2.0)
Other	55.7 (46.1, 65.4)	1.6 (1.0, 2.4)	1.5 ^d (0.9, 2.3)
Frequent e-cigarette use^c			
Race/ethnicity			
NH White	36.3 (32.9, 39.7)	1 (Ref)	1 (Ref)
NH Black	13.5 (8.4, 18.6)	0.3 (0.2, 0.4)	0.2 ^d (0.1, 0.4)
Hispanic/Latino	21.2 (17.3, 25.1)	0.5 (0.4, 0.6)	0.5 ^d (0.4, 0.7)
Other	28.3 (19.7, 36.9)	0.7 (0.4, 1.1)	0.8 ^d (0.5, 1.3)
Flavored e-cigarette use^e			
Race/ethnicity			
NH White	75.0 (72.5, 77.6)	1 (Ref)	1 (Ref)
NH Black	43.9 (37.1, 50.7)	0.3 (0.2, 0.3)	0.4 ^f (0.3, 0.5)
Hispanic/Latino	63.0 (58.6, 67.3)	0.6 (0.5, 0.7)	0.7 ^f (0.5, 0.9)
Other	66.8 (57.5, 76.1)	0.7 (0.4, 1.0)	0.8 ^f (0.5, 1.2)

Note. AOR = adjusted odds ratio; CI = confidence interval; NH = non-Hispanic; OR = odds ratio. The sample size was n = 3628.

^aDual use (co-use of e-cigarettes and cigarettes or other tobacco products) was the dependent variable (yes vs no).

^bAdjusted by age, gender, and tobacco use by a household member.

^cOccasional use (1–5 days in the past 30 days) and frequent use (≥ 20 days in the past 30 days) were the dependent variables (yes vs no) and analyzed separately.

^dAdjusted by age, gender, current cigarette smoking, current other tobacco use, and tobacco use by a household member.

^eFlavored e-cigarette use (tasted like menthol [mint], alcohol [wine, cognac], candy, fruit, chocolate, or any other flavors) was the dependent variable (yes vs no). Flavored e-cigarette use was correlated with the number of days using e-cigarettes in the past 30 d. In 2019, the proportions of flavored e-cigarette use among current e-cigarette users reporting e-cigarettes 1–2 days, 3–5 days, 6–9 days, 10–19 days, 20–29 days, and all 30 days were 54.2%, 66.4%, 71.6%, 76.0%, 79.6%, and 84.1%, respectively.

^fAdjusted by age, gender, current cigarette smoking, current other tobacco use, tobacco use by a household member, and the number of days using e-cigarettes in the past 30 days.

during 2014 to 2019. Prevalence estimates with similar definitions have been reported in other surveillance reports.² Furthermore, some confounders that may affect the outcomes, such as peer use, household income, parental education, and socioeconomic status, were not asked in the NYTS.

Public Health Implications

The latest statistics from the 2020 NYTS data showed a drop in e-cigarette use among US youths, but the prevalence is still at an unacceptably high level, with about 1 in 5 high-school students (~3 million) and 1 in 20 middle-school

students (~550 000) reporting current e-cigarette use.³⁵ This study identified significant disparities in youth e-cigarette use across racial/ethnic groups. Non-Hispanic White youths were more likely to report frequent e-cigarette use and use of flavored products, while non-Hispanic Black

adolescents were more likely to report occasional use of e-cigarettes and dual use of e-cigarettes and cigarettes or other tobacco products than their peers. Tailored, culturally relevant messaging and interventions may help address the racial/ethnic disparities in e-cigarette use. *AJPH*

ABOUT THE AUTHORS

Hongying Dai is with the Department of Biostatistics, College of Public Health, University of Nebraska Medical Center, Omaha. Athena K. Ramos is with the Department of Health Promotion, College of Public Health, University of Nebraska Medical Center. Babalola Faseru is with the Department of Population Health, University of Kansas Medical Center, Kansas City. Jennie L. Hill is with the Department of Epidemiology, College of Public Health, University of Nebraska Medical Center. Steven Y. Sussman is with the Department of Preventive Medicine, University of Southern California Keck School of Medicine, Los Angeles.

CORRESPONDENCE

Correspondence should be sent to Hongying Dai, 84375 Nebraska Medical Center, College of Public Health, University of Nebraska Medical Center, Omaha, NE 68198-4375 (e-mail: daisy.dai@unmc.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

H. Dai and S.Y. Sussman conceptualized and designed the study and drafted the article. All authors acquired, analyzed, or interpreted the data and critically revised the article for important intellectual content. H. Dai performed statistical analysis and obtained funding.

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Note. The content is solely the responsibility of the authors and does not necessarily represent the official views of the National Institutes of Health or the Food and Drug Administration.

CONFLICTS OF INTEREST

The authors report no potential conflicts of interest.

HUMAN PARTICIPANT PROTECTION

Institutional board review was not required for this study, as it was a secondary analysis of de-identified data.

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Our Communities Our Sexual Health ***Awareness and Prevention for African Americans***

Edited By: Madeline Sutton, MD, MPH;
Jo A. Valentine, MSW; and
William C. Jenkins, PhD, MS, MPH

This groundbreaking book provides a comprehensive historical perspective of the disproportionate burden of HIV and other sexually transmitted infections (STIs) among African Americans. Chapters that follow explore the context of HIV and STIs in African American communities and include discussions of sexuality and the roles of faith and spirituality in HIV and STI prevention efforts. Additional chapters provide insight into strategies, e.g., HIV testing, condom distribution and marketing campaigns, parent-child communication, effective clinical care and support, and partnerships, for addressing HIV and other STI-related health disparities within these communities. The book is a valuable resource for practitioners, scholars, clinicians, educators, providers, policy makers and students.



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Sexual Orientation and Gender Identity Data Collection at US Health Centers: Impact of City-Level Structural Stigma in 2018

Anthony N. Almazan, BA, Dana King, ALM, Chris Grasso, MPH, Sean Cahill, PhD, Micah Lattanner, PhD, Mark L. Hatzenbuehler, PhD, and Alex S. Keuroghlian, MD, MPH

Objectives. To examine the relationship between city-level structural stigma pertaining to sexual orientation and gender identity (SOGI) and completeness of patient SOGI data collection at US federally qualified health centers (FQHCs).

Methods. We used the Human Rights Campaign's Municipal Equality Index to quantify city-level structural stigma against sexual and gender minority people in 506 US cities across 49 states. We ascertained the completeness of SOGI data collection at FQHCs from the 2018 Uniform Data System, which describes FQHC patient demographics and service utilization. We included FQHCs in cities captured by the structural stigma index in multinomial generalized linear mixed models to examine the relationship between city-level structural stigma and SOGI data completeness.

Results. FQHCs in cities with more protective sexual orientation nondiscrimination policies reported more complete patient sexual orientation data (adjusted odds ratio [AOR] = 1.6; 95% confidence interval [CI] = 1.2, 2.1). This association was also found for gender identity nondiscrimination policies and gender identity data collection (AOR = 1.7; 95% CI = 1.3, 2.2).

Conclusions. Municipal sexual and gender minority nondiscrimination laws are associated with social and municipal environments that facilitate patient SOGI data collection. (*Am J Public Health.* 2021;111(11):2059–2063. <https://doi.org/10.2105/AJPH.2021.306414>)

Collection of patient sexual orientation and gender identity (SOGI) data is essential for improving health care access, quality, and outcomes for sexual and gender minority (SGM) populations.^{1,2} Given these potential benefits, the US Health Resources and Services Administration Bureau of Primary Health Care requires SOGI data reporting from Health Center Program grantees.³ Federally qualified health centers (FQHCs) provide comprehensive primary care for more than 28 million patients in medically

underserved areas.⁴ Optimizing SOGI data collection at FQHCs is therefore critical for addressing the health needs of the most vulnerable SGM communities.⁵

SOGI data collection approaches across FQHCs are highly variable. These include self-report at intake or registration, report by caregivers, and collection by frontline staff or providers. Although each FQHC is required to report SOGI data annually in aggregate, the lack of consistent mechanisms to ensure SOGI data collection remains a barrier to data

completeness.^{5,6} In an examination of first-year SOGI data reporting by FQHCs, sexual orientation (SO) data were missing for 77% of patients, and gender identity (GI) data were missing for 62%.⁶ Anti-SGM stigma has been hypothesized to be a key driver of limited SOGI data collection. Although the mechanisms underlying this relationship are not fully understood, it may be mediated in part by a lack of training on SOGI data collection, health care staff discomfort, patient nondisclosure, and other confounders.^{6,7}

Studies on the public health implications of stigma have increasingly focused on structural stigma, defined as societal conditions, cultural norms, and institutional policies that adversely affect stigmatized populations.⁸ This research has demonstrated that structural stigma is associated with identity concealment, self-stigma, perceptions of discrimination, poor health behaviors, psychological distress, and psychiatric morbidity across numerous stigmatized populations, including SGM people.⁹ Little is known, however, about the empirical relationship between structural stigma and SOGI data collection.

We conducted the first, to our knowledge, investigation of associations between structural stigma and patient SOGI data collection. We used an existing structural stigma index to examine the relationship between completeness of SOGI data collection at FQHCs and structural stigma related to SOGI.

METHODS

The primary predictors were structural stigma scores. We obtained our structural stigma index from the Human Rights Campaign's 2018 Municipal Equality Index (HRC-MEI),¹⁰ which scores SGM inclusivity for 506 US cities, including the 50 state capitals, the 200 most populated cities, the 5 most populated cities in each state, cities that are home to each state's 2 largest public universities, 75 cities with the highest proportions of same-sex couples, and 98 additional cities selected for inclusion by the HRC and affiliates.

The HRC-MEI measures SGM inclusivity across 5 domains in US cities: nondiscrimination laws (SGM protections in employment and housing), municipality as employer (policies on SGM city employees), municipal services

(inclusiveness of city services), law enforcement (SGM police task forces), and leadership on SGM equality (SGM-affirming public statements and legislation introduced by city leadership). When possible, the HRC-MEI provides separate scores for performance on each item as it relates to SO and GI. Scoring criteria are available.¹⁰

For each city, we generated 5 structural stigma scores (1 for each HRC-MEI domain) by calculating z scores for each city's performance relative to all cities in the index. We calculated separate domain scores for SO and GI. Higher scores represent greater SGM inclusiveness and thus less anti-SGM structural stigma.

The primary outcome was SOGI data completeness, operationalized as the percentage of patients without missing SOGI data at each FQHC. Incomplete data included patients who either declined to answer SOGI questions or were not asked these questions. We obtained FQHC data from the Bureau of Primary Health Care's 2018 Uniform Data System (UDS), an annual data set on FQHC patient demographics and service utilization. The 2018 UDS included data from 1362 FQHCs.

We restricted our analysis of the 2018 UDS data to the 447 FQHCs that could be matched by address to 1 of the cities represented in our structural stigma index. FQHCs in 49 states were represented. City-level structural stigma scores for Hawaii and the District of Columbia were not included in the HRC-MEI.

We conducted our analysis in SAS version 9.4 (SAS Institute, Cary, NC). Two multinomial generalized linear mixed models, 1 each for SO and GI data completeness, tested associations between structural stigma scores and percentages of patients with SO and GI

data (grouped by quartile because of nonnormal distributions and to increase statistical efficiency), with cities as a random intercept. Each model included all 5 structural stigma domain scores and was adjusted for FQHC patient population size (< 10 000, 10 000–19 999, 20 000–29 999, 30 000–49 999, \geq 50 000), percentages of patients who were younger than 18 years, were racial/ethnic minorities, were uninsured, or had income at or below the federal poverty level (based on the US Department of Health and Human Services 2018 federal poverty guidelines). We performed Bonferroni correction for 10 tests, with a significance threshold of *P* of less than .005.

RESULTS

We present the results in [Table 1](#).

The median percentage of patients with complete GI data was 90.0% (interquartile range [IQR] = 44.9%) compared with 67.3% (IQR = 55.3%) for complete SO data. In the final multivariable models, nondiscrimination laws were significantly associated with SO and GI data completeness. The odds of an FQHC being in a higher quartile of data completeness increased with each 1-point increase in structural stigma z score for nondiscrimination laws. This was true for both SO nondiscrimination laws (adjusted odds ratio [AOR] = 1.6; 95% confidence interval [CI] = 1.2, 2.1) and GI nondiscrimination laws (AOR = 1.7; 95% CI = 1.3, 2.2).

Most sociodemographic covariates were not significantly associated with SOGI data completeness, including FQHC size, percentage of racial/ethnic minority patients, percentage below the federal poverty level, and percentage uninsured. The percentage of patients younger than 18 years was significantly associated with

TABLE 1— Associations Between City-Level Structural Stigma Scores and the Percentage of Patients With Complete Sexual Orientation and Gender Identity Data Collection at Federally Qualified Health Centers (FQHCs): United States, 2018

	Quartile 1, % Range, No., or Mean (SD)	Quartile 2, % Range, No., or Mean (SD)	Quartile 3, % Range, No., or Mean (SD)	Quartile 4, % Range, No., or Mean (SD)	Multivariable Model for Higher Quartile of Data Completeness, AOR (95% CI)
Sexual orientation					
Range	0.0–37.1	37.2–69.5	69.5–92.6	92.6–100.0	
Sample size	112	120	111	104	
Structural stigma domain					
Nondiscrimination laws	0.4 (1.0)	0.6 (1.0)	0.7 (1.0)	0.8 (1.0)	1.6 (1.2, 2.1) ^a
Municipality as employer	0.6 (0.8)	0.6 (0.8)	0.5 (0.8)	0.6 (0.7)	0.9 (0.7, 1.2)
Municipal services	0.6 (0.9)	0.8 (0.9)	0.8 (0.9)	0.7 (0.9)	0.9 (0.7, 1.2)
Law enforcement	0.6 (0.9)	0.7 (0.9)	0.8 (0.7)	0.8 (0.8)	1.4 (1.0, 2.0)
Leadership on SGM equality	0.7 (0.9)	0.7 (0.8)	0.6 (0.9)	0.7 (0.9)	0.7 (0.5, 1.0)
Gender identity					
Range	0.0–55.3	55.4–90.8	90.9–99.8	99.9–100.0	
Sample size	113	117	104	113	
Structural stigma domain					
Nondiscrimination laws	0.4 (1.1)	0.3 (1.1)	0.7 (1.1)	0.8 (1.1)	1.7 (1.3, 2.2) ^a
Municipality as employer	0.7 (0.9)	0.6 (1.0)	0.8 (0.9)	0.7 (1.0)	0.7 (0.5, 1.0)
Municipal services	0.7 (1.0)	0.8 (1.0)	0.9 (1.0)	0.7 (1.0)	1.0 (0.7, 1.4)
Law enforcement	0.7 (0.8)	0.5 (1.0)	0.8 (0.7)	0.7 (0.8)	1.1 (0.8, 1.5)
Leadership on SGM equality	0.7 (0.9)	0.6 (0.9)	0.8 (0.8)	0.7 (0.9)	0.9 (0.6, 1.4)

Note. AOR = adjusted odds ratio; CI = confidence interval; SGM = sexual and gender minority. All models were adjusted for FQHC size and the percentage of patients at each FQHC who were younger than 18 years, identified as racial/ethnic minorities, were uninsured, and had incomes at or below the federal poverty level.

^aSignificant based on Bonferroni correction for 10 tests.

SO data completeness (AOR = 16.8; 95% CI = 3.5, 80.7) and GI data completeness (AOR = 10.6; 95% CI = 2.5, 45.1), but these estimates were unstable and should be interpreted with caution.

DISCUSSION

This is the first study, to our knowledge, to demonstrate an empirical relationship between structural stigma and patient data collection. FQHCs in cities with more protective SGM nondiscrimination laws reported more complete SOGI patient data than did FQHCs in

cities with less protective nondiscrimination laws. These findings support the hypothesis that anti-SGM stigma is associated with limited SOGI data collection.⁶ Our focus on city-level structural stigma adds to the existing stigma literature, as scholars have hypothesized that proximal (e.g., city-level) forms of structural stigma may have greater influences on health than distal (e.g., state-level) forms.⁹

Notably, the only structural stigma domain significantly associated with SOGI data collection was nondiscrimination laws. There were no statistically significant

associations between SOGI data completeness and the 4 other structural stigma domains. These findings suggest that enforceable SGM-inclusive laws are more strongly associated with stigma alleviation and SOGI data collection than are the other structural stigma domains, which were largely composed of voluntary practices and symbolic representations of SGM inclusivity.

The completeness of GI data may be substantially higher than the completeness of SO data in this sample for several reasons. First, sex-related data, which may be conflated with GI data, are

collected more consistently for billing purposes. Second, it is possible that frontline staff do not explicitly ask GI questions; they might assume GI based on their own perception of each patient's gender expression. Third, it is possible that health care staff may view GI as more actionably relevant to medical care than SO.

Our study has several limitations. Our structural stigma index is based on the HRC-MEI; therefore, our findings are specific to FQHCs in the HRC-MEI's 506 listed cities and may not be generalizable across all health care systems. The UDS does not distinguish between SOGI data being incomplete because patients were not asked by staff versus patients chose not to disclose, which precluded assessment of mechanisms underlying this study's findings. Furthermore, the UDS only reports aggregate data at the FQHC level, precluding patient-level statistical analysis. The study's cross-sectional design does not allow us to determine causation or directionality of our main finding, which may be bidirectional: SGM nondiscrimination laws may promote SOGI data collection,¹¹ and SOGI data collection may facilitate the implementation of SGM-inclusive policies.¹² Finally, as this is an observational study, it may be subject to unmeasured confounding.

Future investigations on structural stigma and SOGI data collection would benefit from the expansion of structural stigma indices to measure stigma in a more diverse selection of regions, including more rural settings. Furthermore, future studies should thoroughly examine how predictors of differential access to affirming health care may influence SOGI data collection. In this study, percentage of racial and ethnic minority patients was not a significant predictor of SOGI data completeness.

However, our access to only aggregate FQHC-level data may have precluded our identification of significant associations. This limitation reinforces the importance of future patient-level analyses for assessing associations between structural stigma and patient data collection.

PUBLIC HEALTH IMPLICATIONS

In this cross-sectional study, we found that structural stigma scores quantifying the protectiveness of city-level SGM nondiscrimination laws were associated with SOGI data collection completeness at FQHCs. Our findings suggest potential value in enhanced federal oversight to improve mandated collection of SOGI data, particularly in cities with higher levels of anti-SGM structural stigma. These findings also underscore the importance of municipal SGM nondiscrimination laws and affirming social environments, which may mutually reinforce one another, foster robust SOGI data collection, and promote SGM health equity. *AJPH*

ABOUT THE AUTHORS

Anthony N. Almazan is with Harvard Medical School, Boston, MA. Dana King, Chris Grasso, Sean Cahill, and Alex S. Keuroghlian are with The Fenway Institute, Fenway Health, Boston. Micah Lattanner and Mark L. Hatzenbuehler are with the Department of Psychology, Harvard University, Boston.

CORRESPONDENCE

Correspondence should be sent to Alex S. Keuroghlian, The Fenway Institute, 1340 Boylston St, Boston, MA 02215 (e-mail: akeuroghlian@partners.org). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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CONTRIBUTORS

A. N. Almazan led the drafting of the article. A. N. Almazan and D. King contributed to analysis design. A. N. Almazan and A. S. Keuroghlian conceptualized and designed the study. D. King conducted the analyses. M. Lattanner and M. L. Hatzenbuehler contributed to the development of the structural stigma index. A. S. Keuroghlian supervised the implementation and reporting of the study. All authors edited the final article.

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CONFLICTS OF INTEREST

A. S. Keuroghlian will receive royalties as editor of an upcoming McGraw-Hill Education textbook on transgender and gender diverse care. The authors report no other relevant conflicts of interest.

HUMAN PARTICIPANT PROTECTION

This study was reviewed by the Fenway Institute institutional review board and did not constitute human participants research.

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Conducting Health Research with Native American Communities

Edited by Teshia G. Arambula Solomon, PhD and Leslie L. Randall, RN, MPH, BSN



The current research and evaluation of the American Indian and Alaska Native (AIAN) people demonstrates the increased demand for efficiency, accompanied by solid accountability in a time of extremely limited resources. This environment requires proficiency in working with these vulnerable populations in diverse cross-cultural settings. This timely publication is the first of its kind to provide this information to help researchers meet their demands.

This book provides an overview of complex themes as well as a synopsis of essential concepts or techniques in working with Native American tribes and Alaska Native communities. *Conducting Health Research with Native American Communities* will benefit Native people and organizations as well as researchers, students and practitioners.

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Self-Reported Health Indicators in the US Army: Longitudinal Analysis From a Population Surveillance System, 2014–2018

Jessica L. Kegel, MA, Josh B. Kazman, MS, Daniel R. Clifton, PhD, ATC, Patricia A. Deuster, PhD, MPH, and Sarah J. de la Motte, PhD, MPH, ATC

Objectives. To describe health-related behaviors or indicators associated with overall health and well-being using the Global Assessment Tool (GAT), a health behavior and psychosocial questionnaire completed annually by US Army personnel.

Methods. We analyzed GAT responses from 2014 to 2018, consisting of 367 000 to 449 000 respondents per year. We used generalized estimating equations to predict the presence of each health behavior or indicator, aggregated by year and stratified on various demographics.

Results. Key findings included decreases from 2014 to 2018 in risky health behaviors such as hazardous drinking (7.5% decrease) and tobacco use (7.9% decrease), dietary supplement use (5.0% to 10.6% decrease, depending on type), self-reported musculoskeletal injury (5.1% decrease), and pain interference (3.6% decrease). Physical activity, sleep, and nutritional habits largely remained consistent over time.

Conclusions. In the Army, tobacco, alcohol, and risky dietary supplement usage appears to be declining, whereas lifestyle health behaviors have been stable. Whether these trends reflect responses to health education is unknown. The GAT provides useful insights into the health of the Army, which can be leveraged when developing health-related educational programs and policies.

Public Health Implications. Health behaviors that have changed less over time (e.g., nutrition, sleep) may require novel approaches compared with those that changed more (e.g., dietary supplement use, drinking). (*Am J Public Health.* 2021;111(11):2064–2074. <https://doi.org/10.2105/AJPH.2021.306456>)

Data repositories of self-reported health-related behaviors and traits, such as the National Health and Nutrition Examination Survey (NHANES),¹ track secular health trends and determine risk for mobility² and musculoskeletal (MSK) disability.³ These efforts provide a roadmap for evaluating population-level health-related behaviors and indicators. The US Army is particularly

concerned about soldier health and deployability and has multiple surveillance systems to track the health status of the force. In 2017, approximately 60 000 US Army soldiers were unfit for combat, with up to 80% of those unable to deploy because of injuries and illness.⁴ With the 2018 Army directive on deployability and fitness, Army soldiers who do not meet deployment standards, or who become

permanently nondeployable, are unqualified to serve.⁵ Thus, it is important to understand factors that affect health and well-being and contribute to deployability.

One of the largest health behavior surveillance systems for the Army is the Global Assessment Tool (GAT), a self-report survey that assesses emotional, social, family, spiritual, and physical health (on October 2019, the GAT

was renamed Azimuth Check [Army Regulation 350–53]).⁶ Soldier health behaviors and related indicators assessed with the GAT provide insights into population-level trends. The GAT provides similar data as NHANES and select other military studies (a regular cross-sectional population-stratified survey of health behaviors⁷ and a large longitudinal cohort study of mental health⁸) covering multiple health-related behaviors and well-being indicators plausibly associated with military deployability. Specifically, the GAT focuses on sleep, physical activity, and nutrition, which have recently been the focus of Army readiness-related programs.⁹ Alcohol consumption (also addressed in the GAT) may also affect job performance, and excessive use can lead to disciplinary actions and associated occupational terminations.¹⁰ Further factors of interest included in the GAT are dietary supplement use,¹¹ tobacco use,¹² and MSK injury.¹³

Describing Army health behaviors provides valuable information to assist in focusing public health efforts within the military community. We employed longitudinal modeling to examine population-level trends for GAT health indicators during 2014 to 2018. We summarize findings in health-promoting and risky health behaviors or indicators across the Army as a whole, stratified by gender, service component (i.e., active duty, National Guard, Army Reserve), rank, years of service, and military occupational specialty.

METHODS

All Army personnel are required to annually complete the GAT and are subsequently given the option to allow their responses to be used for research

purposes. For this study, data consisted of GAT and demographic records from 2014 to 2018 of US Army soldiers who consented to have their responses used for research, resulting in a dynamic cohort that varies by study year (total $n = 952\,371$; see Table 1 for breakdown by year). On average, each individual in the cohort recorded 2.1 ± 1.2 observations over the period from 2014 to 2018. Most respondents were White, male, enlisted, and active-duty soldiers aged 34 years or younger across all 5 years (Table 1).

GAT data are stored in the Person-Event Data Environment, a secure data enclave maintained by the Army Analytics Group Research Facilitation Laboratory (Monterey, CA), combining large data sets de-identified for research. In this study, demographic and career-related variables were culled from administrative data sets from the Defense Manpower Data Center and merged with a data set of GAT respondents.

GAT data from about 60% of the Army flows to the Person-Event Data Environment; within those data, about 70% of participants consent for their GAT responses to be used for research and were thus accessed for our analyses. We compared this analytic GAT sample with total Army reports¹⁴ and estimate that our sample represents about 40% of the Army. Demographically, our sample is very similar to the total Army (see Table A, available as a supplement to the online version of this article at <http://www.ajph.org>, for comparisons stratified by year and component).

Measures

Key variables are summarized here; more extensive details regarding

survey item wording are provided in the Appendix (available as a supplement to the online version of this article at <http://www.ajph.org>; see Section 1 – Item Wording). To concisely present the results, all health behavior or indicator items were dichotomized or categorized by using public health guidelines where applicable (e.g., fruit and vegetable consumption,¹⁵ activity,¹⁶ and sleep¹⁷), or expert opinion where guidelines were lacking.

Demographic information, including gender, age, race, career path, and service component were accessed through the personnel records provided by the Defense Manpower Data Center.

Regarding physical activity, soldiers were asked about the duration and frequency of vigorous aerobic activity, which was then coded into activity minutes per week and dichotomized based on 150 or more minutes per week. For days per week of resistance training, responses were classified as low (0–1 days/week), moderate (2–4 days/week), or high (≥ 5 days/week). A final question asked about high-intensity interval training (HIIT, defined in the survey as “sets of high intensity exercises with brief rest periods in between”) and was dichotomized based on soldiers reporting any HIIT activity in the past 30 days.

To assess dietary habits, frequency of fruit and vegetable intake, breakfast consumption, and postexercise recovery snacks were analyzed. For fruit and vegetable intake, respondents were dichotomized based on whether they ate 2 or more servings per day or less. For breakfast, soldiers were dichotomized as regularly eating breakfast (6 to 7 times per week) and as not regularly eating breakfast (fewer than 6 times per week). Postexercise recovery fueling

TABLE 1— Sample Demographics of Global Assessment Tool Respondents by Year: United States, 2014–2018

	Year, No. or %				
	2014	2015	2016	2017	2018
Sample	422 044	449 353	395 051	366 586	343 432
Gender					
Male	83.3	82.7	82.5	81.9	81.6
Female	16.7	17.3	17.5	18.1	18.4
Age, y					
17–24	30.7	33.2	35.3	35.7	34.1
25–34	38.9	37.5	36.4	35.9	36.4
35–44	20.6	19.6	18.7	18.6	19.4
≥ 45	9.8	9.7	9.6	9.9	10.2
Race					
American Indian/Alaska Native	0.7	0.7	0.7	0.7	0.7
Asian	3.7	4.0	4.4	4.6	4.7
Black or African American	20.0	20.6	20.3	20.6	19.9
Native Hawaiian or other Pacific Islander	0.7	0.8	0.8	0.8	0.9
White	70.3	69.7	69.7	69.6	70.1
Other or unknown	4.6	4.3	4.2	3.8	3.8
Rank					
Enlisted	81.2	81.3	82.7	82.4	81.8
Officer	18.8	18.7	17.3	17.6	18.2
Service component					
Active duty	61.0	57.5	51.2	50.0	51.5
Reserve	15.8	17.2	20.4	21.1	21.8
National Guard	23.2	25.3	28.4	28.8	26.7

responses were categorized as frequently engaging in recovery fueling (“often” or “most of the time”) and as never or infrequently engaging in recovery fueling (“only sometimes” or less).

Dietary supplement use was dichotomized based on whether soldiers reported using products at least once a month in the previous year. Respondents who affirmatively responded to taking dietary supplements were then asked about dietary supplement use in the following categories: health-promoting, vitamin D, fish oil, sport protein powders, weight loss, and performance-enhancing or body-building dietary supplements.

Sleep duration was assessed with 2 items that asked weekday and weekend sleep hours. Responses were combined and dichotomized at 6 hours or less per average day versus more than 6 hours. Sleep quality was assessed using the 2-item Pittsburgh Insomnia Rating Scale, which evaluates sleep satisfaction and difficulties relating to lack of energy because of poor sleep within the last week.¹⁸ A score greater than 2 was used to classify soldiers as at risk for insomnia.

Frequency of tobacco use was categorized as regular tobacco users (using tobacco for the past 3 months or more) and nontobacco users

(not using tobacco within the past 3 months).

The Alcohol Use Disorders Identification Test (AUDIT), a 2-item screen (assessing frequency and quantity of alcohol consumption), was used.¹⁹ Soldiers categorized as hazardous drinkers had AUDIT scores greater than or equal to 4, and soldiers categorized as non-hazardous drinkers had AUDIT scores less than 4.

Self-reported limited duty within the past year because of a MSK injury was assessed to indicate whether soldiers were deemed medically nonready (MSK-MNR). Responses were dichotomized as “yes” or “no” based on

whether they responded affirmative to having sustained a duty-limiting MSK injury within 12 months. Pain interference in usual activities was assessed along a scale of 1 (“hardly notice”) to 10 (“as bad as it could be, nothing else matters”). Responses were dichotomized as a rating of 5 (“interrupts some activities”) or higher and 4 (“distracts me, can do usual activities”) or below.

Analyses

For descriptive analyses, responses were aggregated by calendar year to examine overall percentage changes in prevalence over the 5-year period from 2014 to 2018. Absolute and average rates of change were calculated from 2014 to 2018. Group differences were then assessed based upon demographics (i.e., gender, age) and occupational variables (i.e., years of service, rank, component, military occupational specialty). Military occupational specialty was categorized as combat support, cyber and military intelligence, engineering and maintenance, front-line, or service support (breakdown provided in the online Appendix, Section 2 – MOS Categories) to obtain an analysis of different career fields. Overall prevalence percentages are described using frequencies of respondents who reported engaging in a certain behavior in each survey year.

For statistical analyses, longitudinal trends and person covariates were modeled at the individual level using generalized estimating equations to predict the presence of each health behavior or indicator, with logit link and binomial distribution and an autoregressive covariance structure accounting for responses from the same respondent in different years (SAS

Enterprise Guide version 7.1, PROC GENMOD; SAS Institute, Cary, NC). Only resistance training was modeled differently to retain its 3 categories and capture overtraining; it was modeled as an ordinal variable by using the cumulative probability of engaging in higher levels of training and with an independence covariance structure. Because there were only 5 time points and no a priori reason to suspect linear changes, “year” was modeled as a categorical variable indicating odds ratios (ORs) for the presence of behaviors or indicators relative to the baseline year (2014).

We executed 2 types of models for each outcome: a time-only longitudinal model, which just used survey year as a predictor, and a covariate-adjusted longitudinal model, which used survey year and demographic and occupational predictors (OR referred to as adjusted OR [AOR]). Results from the majority of time-only models closely mirrored those of the covariate-adjusted models, and findings from both models can be found in the tables. Most results presented here focus on health behaviors in 2018 compared with 2014 in covariate-adjusted models. Finally, time-only models were also conducted within stratified samples, with strata consisting of various demographic and occupational variables. Results from these stratified models are presented as ORs and our interpretation focuses on effect sizes. Notable results from the stratified analyses are highlighted here.

RESULTS

With regard to physical activity, respondents had greater odds of engaging in at least 150 minutes

of vigorous activity in 2018 compared with 2014 (Table 2: 53% vs 52% of respondents; Table 3: AOR = 1.11). Resistance training more than 5 days per week increased overall by 2% in 2018 compared with 2014 (Table 3: AOR = 1.18). Respondents had lower odds of participating in HIIT programs at least once per week in 2018 compared with 2014 (Table 2: 34% vs 38%; Table B [available as a supplement to the online version of this article at <http://www.ajph.org>] AOR = 0.83).

In stratified analyses comparing prevalence rates in 2014 to 2018, the largest exceptions to these trends were based on service duration: soldiers with 6 to 10 years of service had a greater increase in reported vigorous activity (Table C [available as a supplement to the online version of this article at <http://www.ajph.org>]: 7% increase; OR = 1.28) and resistance training (6% increase; OR = 1.39) from 2014 to 2018. However, there was a smaller change in HIIT participation among soldiers with 6 to 10 (Table C: 0% change; OR = 0.94) and 11 to 30 years of service (2% increase; OR = 1.02). There were no differences for those who had served 0 to 5 years.

With regard to dietary habits, over the 5-year period, fruit consumption of 2 or more servings per day ranged from 36% in 2014 and 2018 to a high of 39% in 2016 (Table 3: 2018 vs 2014 AOR = 0.89), and vegetable consumption of 2 or more servings per day ranged from 45% in 2014 to a high of 48% in 2016 (Table 3: 2018 vs 2014 AOR = 0.98). Regular breakfast consumption (6–7 days/week) was higher in 2016 (55%) but decreased to 53% in 2018 compared with 2014 (Table B: 2018 vs 2014 AOR = 1.04).

TABLE 2— Trends in Healthy and Risky Behaviors Across the US Army From 2014 to 2018

	% Reported by Year					Difference Between 2018 and 2014	Average Difference
	2014	2015	2016	2017	2018		
Physical activity							
High vigorous (> 150 min/wk)	51.7	52.4	53.6	52.6	53.4	1.7	0.4
Moderate resistance (2–4 d/wk)	46.7	46.1	45.4	45.9	47.1	0.4	0.1
High resistance (≥ 5 d/wk)	30.7	31.7	33.5	33.2	32.9	2.2	0.4
HIIT (≥ 1 d/wk)	38.4	35.3	34.9	33.6	33.6	−4.9	−1.0
Dietary habits							
Fruit (≥ 2 servings/d)	36.3	37.5	38.7	37.8	36.0	−0.3	−0.1
Vegetable (≥ 2 servings/d)	45.3	46.5	47.8	47.3	46.6	1.4	0.3
Breakfast (6–7 d/wk)	48.9	51.7	55.0	54.2	52.5	3.6	0.7
Recovery fueling (≥ often)	52.1	52.9	53.7	54.0	53.2	1.1	0.2
Dietary supplement use							
Health promoting	24.1	16.0	14.5	13.7	13.6	−10.6	−2.1
Vitamin D	19.4	12.4	11.2	10.6	10.5	−9.0	−1.8
Fish oil	18.0	12.0	10.7	10.1	9.8	−8.2	−1.6
Protein	20.2	13.9	12.9	12.5	12.3	−7.9	−1.6
Weight loss	9.9	6.4	5.5	5.2	4.9	−5.0	−1.0
Performance enhancement	12.2	8.4	7.4	7.0	6.5	−5.7	−1.1
Sleep							
Poor sleep quality (PIRS > 2)	43.4	42.5	42.7	41.8	39.9	−3.5	−0.7
Short sleep (≤ 6 h/d)	35.1	33.8	33.3	33.0	32.1	−2.9	−0.6
Risky health behaviors							
Hazardous drinking (AUDIT > 3)	26.7	24.5	21.8	20.5	19.2	−7.5	−1.5
Tobacco	24.5	21.1	18.8	17.5	16.6	−7.9	−1.6
Medical							
Musculoskeletal Injury	34.9	31.4	29.7	29.1	29.8	−5.1	−1.0
Pain interference	21.0	19.3	18.2	17.6	17.5	−3.6	−0.7

Note. AUDIT = Alcohol Use Disorder Identification Test;¹⁹ HIIT = high-intensity interval training; PIRS = 2-item Pittsburgh Insomnia Rating Scale.¹⁸ Exact item wording is provided in Appendix (available as a supplement to the online version of this article at <http://www.ajph.org>; see Section 1—Item Wording).

Approximately 53% of respondents frequently consumed postexercise recovery fueling in 2018 compared with 52% in 2014 (Table B: AOR = 1.07).

Stratified unadjusted analyses suggested some differences in these trends by gender and service duration. Women (Table C: 3% decrease; OR = 0.85) and soldiers with 6 to 10 years of service (5% decrease; OR = 0.80) reported a greater decrease in fruit consumption compared with

the increase in the full sample. Soldiers holding a cyber and military intelligence military occupational specialty reported a decrease in regular fruit (Table C: 4% decrease; OR = 0.79) and breakfast (2% decrease; OR = 0.87) consumption compared with the full sample.

A 5% to 11% decrease was reported in all categories of dietary supplement use (health-promoting, vitamin D, fish oil, protein, weight-loss, and performance-enhancing) in 2018 compared

with 2014 (Table B: AOR ranging from 0.52 to 0.61). Similar trends were seen among the stratified groups compared with the full sample.

A decrease was reported in poor sleep quality (Table 2: 4%; Table 4: AOR = 0.86) and short sleep duration (i.e., 6 hours or less; Table 2: 3%; Table 4: AOR = 0.88) in 2018 compared with 2014. Similar trends were seen among the stratified groups compared with the full sample.

TABLE 3— Trends and Group Differences of Physical Activity and Food and Dietary Supplement Consumption Derived From Generalized Estimating Equations Among Respondents to the Global Assessment Tool: United States, 2014–2018

	Physical Activity, AOR (95% CI)		Dietary Habits, AOR (95% CI)		Dietary Supplements, AOR (95% CI)	
	Vigorous	Resistance	Fruit	Vegetables	Weight-Loss	Performance
Time-only model^a						
Survey Year (Ref: 2014)						
2015	1.02 (1.02, 1.03)	1.04 (1.03, 1.04)	1.05 (1.04, 1.05)	1.05 (1.04, 1.05)	0.61 (0.60, 0.62)	0.64 (0.63, 0.65)
2016	1.07 (1.06, 1.08)	1.12 (1.11, 1.13)	1.09 (1.08, 1.10)	1.09 (1.09, 1.10)	0.52 (0.51, 0.53)	0.57 (0.56, 0.58)
2017	1.04 (1.03, 1.05)	1.12 (1.11, 1.13)	1.05 (1.04, 1.06)	1.07 (1.07, 1.08)	0.50 (0.49, 0.51)	0.54 (0.54, 0.55)
2018	1.05 (1.04, 1.06)	1.13 (1.12, 1.14)	0.96 (0.95, 0.96)	1.03 (1.02, 1.04)	0.48 (0.47, 0.48)	0.51 (0.50, 0.52)
Covariate-adjusted model^b						
Survey year (Ref: 2014)						
2015	1.04 (1.03, 1.05)	1.04 (1.03, 1.05)	1.03 (1.02, 1.04)	1.04 (1.03, 1.05)	0.65 (0.63, 0.66)	0.67 (0.66, 0.69)
2016	1.11 (1.10, 1.12)	1.16 (1.15, 1.17)	1.13 (1.11, 1.14)	1.14 (1.13, 1.15)	0.54 (0.52, 0.55)	0.58 (0.57, 0.59)
2017	1.08 (1.07, 1.10)	1.17 (1.16, 1.19)	1.05 (1.04, 1.06)	1.09 (1.08, 1.10)	0.53 (0.52, 0.54)	0.57 (0.56, 0.58)
2018	1.11 (1.10, 1.13)	1.18 (1.17, 1.20)	0.89 (0.88, 0.91)	0.98 (0.97, 0.99)	0.52 (0.51, 0.53)	0.56 (0.55, 0.57)
Female gender (Ref: male)	0.67 (0.67, 0.68)	0.64 (0.63, 0.65)	1.33 (1.31, 1.35)	1.29 (1.27, 1.30)	1.18 (1.15, 1.21)	0.46 (0.45, 0.48)
Age, y (Ref: 25–34)						
17–24	1.33 (1.31, 1.34)	1.33 (1.32, 1.34)	1.09 (1.08, 1.10)	0.88 (0.87, 0.89)	0.69 (0.68, 0.71)	0.78 (0.77, 0.79)
35–44	0.81 (0.80, 0.82)	0.72 (0.72, 0.73)	0.88 (0.87, 0.89)	0.92 (0.91, 0.93)	1.08 (1.05, 1.10)	0.80 (0.79, 0.82)
≥ 45	0.71 (0.70, 0.72)	0.60 (0.59, 0.61)	0.83 (0.82, 0.85)	0.89 (0.87, 0.90)	0.82 (0.79, 0.85)	0.44 (0.43, 0.46)
Officer rank (Ref: enlisted)	0.84 (0.83, 0.85)	0.97 (0.96, 0.98)	1.44 (1.42, 1.46)	1.57 (1.55, 1.59)	0.69 (0.67, 0.71)	0.96 (0.94, 0.99)
Service component (Ref: active duty)						
National Guard	0.49 (0.48, 0.49)	0.61 (0.60, 0.62)	0.96 (0.95, 0.97)	1.00 (0.99, 1.01)	0.66 (0.65, 0.68)	0.74 (0.72, 0.75)
Reserve	0.60 (0.59, 0.62)	0.74 (0.73, 0.76)	1.08 (1.06, 1.10)	1.05 (1.03, 1.07)	0.79 (0.76, 0.82)	0.80 (0.77, 0.83)
Length of service, y (Ref: 0–5)						
6–10	0.85 (0.84, 0.86)	0.84 (0.83, 0.85)	0.92 (0.91, 0.93)	0.95 (0.94, 0.96)	0.99 (0.97, 1.02)	0.97 (0.95, 0.99)
11–30	0.78 (0.77, 0.80)	0.75 (0.74, 0.77)	0.98 (0.96, 1.01)	1.00 (0.98, 1.03)	0.90 (0.86, 0.94)	0.74 (0.71, 0.77)
Occupational specialty (Ref: combat support)						
C&M intelligence	0.90 (0.89, 0.92)	1.01 (0.99, 1.03)	1.08 (1.06, 1.10)	1.22 (1.19, 1.24)	1.03 (0.99, 1.07)	1.16 (1.12, 1.2)
E&M	1.10 (1.08, 1.11)	1.05 (1.04, 1.07)	0.97 (0.96, 0.99)	0.98 (0.97, 0.99)	0.93 (0.91, 0.96)	0.96 (0.93, 0.98)
Frontline	1.31 (1.30, 1.33)	1.34 (1.32, 1.35)	1.11 (1.09, 1.12)	1.15 (1.13, 1.16)	1.02 (1.00, 1.04)	1.19 (1.17, 1.21)
Service support	0.87 (0.86, 0.88)	0.93 (0.92, 0.94)	1.18 (1.16, 1.19)	1.21 (1.20, 1.23)	0.86 (0.84, 0.88)	0.91 (0.89, 0.94)

Note. AOR = adjusted odds ratio; C&M = cyber and military; CI = confidence interval; E&M = engineering and maintenance. Separate models were conducted for each column, adjusting for all the demographics listed on the rows.

^aModel only used year as a predictor (1 976 000 observations from 952 000 respondents).

^bModel used both year and demographic and occupational covariates listed in table as predictors (1 227 000 observations from 583 000 respondents).

Reports of hazardous drinking were lower in 2018 (19%) compared with 2014 (Table 2: 27%; Table 4: AOR = 0.72). Tobacco use decreased by

8% during this 5-year period (Table 4: AOR = 0.73).

In stratified analyses, soldiers with 11 to 30 years of service (Table B: 4%;

OR = 0.90) reported a smaller decrease in tobacco use than the full sample.

Overall, reported MSK-MNR was 30% in 2018 compared with 35% in 2014

TABLE 4— Trends and Group Differences for Sleep, Substance Use, and Musculoskeletal Outcomes Derived From Generalized Estimating Equations Among Army Respondents to the Global Assessment Tool: United States, 2014–2018

	Sleep, AOR (95% CI)		Risky Health Behaviors, AOR (95% CI)		Medical Outcomes, AOR (95% CI)	
	Poor Quality	Short Duration	Hazardous Drinking	Tobacco Use	MSK-MNR	Pain
Time-only model^a						
Survey year (Ref: 2014)						
2015	0.98 (0.97, 0.99)	0.96 (0.95, 0.97)	0.90 (0.89, 0.91)	0.87 (0.87, 0.88)	0.88 (0.87, 0.89)	0.93 (0.92, 0.94)
2016	0.99 (0.98, 0.99)	0.94 (0.94, 0.95)	0.78 (0.78, 0.79)	0.79 (0.78, 0.79)	0.82 (0.82, 0.83)	0.88 (0.87, 0.89)
2017	0.95 (0.94, 0.95)	0.92 (0.91, 0.93)	0.73 (0.72, 0.74)	0.74 (0.74, 0.75)	0.81 (0.81, 0.82)	0.85 (0.84, 0.86)
2018	0.87 (0.86, 0.88)	0.89 (0.88, 0.90)	0.69 (0.68, 0.70)	0.71 (0.71, 0.72)	0.86 (0.85, 0.87)	0.85 (0.84, 0.86)
Covariate-adjusted model^b						
Survey year (Ref: 2014)						
2015	0.93 (0.92, 0.94)	0.91 (0.90, 0.92)	0.91 (0.90, 0.92)	0.87 (0.87, 0.88)	0.92 (0.91, 0.93)	0.95 (0.93, 0.96)
2016	1.00 (0.99, 1.01)	0.95 (0.94, 0.96)	0.76 (0.75, 0.77)	0.75 (0.74, 0.76)	0.84 (0.83, 0.85)	0.90 (0.88, 0.91)
2017	0.95 (0.94, 0.96)	0.91 (0.90, 0.92)	0.73 (0.72, 0.73)	0.73 (0.73, 0.74)	0.84 (0.84, 0.85)	0.87 (0.86, 0.88)
2018	0.86 (0.85, 0.87)	0.88 (0.87, 0.89)	0.72 (0.71, 0.72)	0.73 (0.72, 0.74)	0.88 (0.87, 0.89)	0.85 (0.84, 0.86)
Female gender (Ref: male)	1.45 (1.43, 1.47)	1.09 (1.08, 1.11)	0.95 (0.94, 0.96)	0.43 (0.42, 0.43)	1.47 (1.45, 1.49)	1.43 (1.41, 1.45)
Age, y (Ref: 25–34)						
17–24	1.10 (1.09, 1.11)	0.88 (0.87, 0.89)	0.87 (0.86, 0.88)	0.81 (0.80, 0.82)	0.62 (0.61, 0.62)	0.66 (0.65, 0.67)
35–44	1.14 (1.13, 1.16)	1.27 (1.26, 1.29)	0.85 (0.84, 0.86)	1.04 (1.03, 1.06)	1.69 (1.67, 1.71)	1.74 (1.72, 1.77)
≥ 45	1.02 (1.00, 1.04)	1.31 (1.29, 1.34)	0.67 (0.66, 0.69)	0.95 (0.93, 0.97)	2.14 (2.11, 2.18)	2.12 (2.08, 2.16)
Officer rank (Ref: enlisted)	0.64 (0.63, 0.64)	0.44 (0.43, 0.45)	1.26 (1.24, 1.28)	0.38 (0.37, 0.39)	0.58 (0.57, 0.59)	0.48 (0.47, 0.49)
Service component (Ref: active duty)						
National Guard	0.71 (0.70, 0.71)	0.66 (0.65, 0.67)	1.00 (0.99, 1.02)	0.83 (0.82, 0.84)	0.41 (0.41, 0.42)	0.50 (0.49, 0.51)
Reserve	0.71 (0.70, 0.73)	0.76 (0.75, 0.78)	0.92 (0.90, 0.94)	0.80 (0.78, 0.82)	0.45 (0.44, 0.46)	0.56 (0.55, 0.57)
Length of service, y (Ref: 0–5)						
6–10	1.06 (1.05, 1.07)	1.07 (1.06, 1.08)	1.06 (1.04, 1.07)	1.04 (1.03, 1.06)	1.12 (1.11, 1.13)	1.18 (1.17, 1.20)
11–30	1.02 (1.00, 1.04)	0.98 (0.96, 1.01)	1.06 (1.04, 1.09)	0.91 (0.88, 0.93)	1.12 (1.10, 1.15)	1.19 (1.16, 1.20)
Occupational specialty (Ref: combat support)						
C&M intelligence	0.95 (0.93, 0.96)	0.76 (0.74, 0.77)	1.08 (1.05, 1.10)	0.77 (0.75, 0.8)	1.10 (1.08, 1.13)	0.93 (0.90, 0.95)
E&M	1.00 (0.99, 1.02)	1.08 (1.06, 1.09)	1.13 (1.12, 1.15)	1.30 (1.28, 1.32)	0.99 (0.97, 1.00)	1.05 (1.04, 1.07)
Frontline	0.97 (0.96, 0.98)	1.03 (1.02, 1.05)	1.22 (1.20, 1.23)	1.33 (1.31, 1.35)	0.88 (0.87, 0.89)	0.97 (0.96, 0.98)
Service support	0.92 (0.90, 0.93)	0.84 (0.83, 0.86)	0.91 (0.90, 0.93)	0.75 (0.74, 0.76)	1.00 (0.99, 1.01)	0.79 (0.78, 0.80)

Note. AOR = adjusted odds ratio; C&M = cyber and military; CI = confidence interval; E&M = engineering and maintenance; MSK-MNR = musculoskeletal injury–medically nonready. Separate models were conducted for each column, adjusting for all the demographics listed on the rows.

^aModel only used year as a predictor (1 976 000 observations from 952 000 respondents).

^bModel used both year and demographic and occupational covariates listed in table as predictors (1 227 000 observations from 583 000 respondents).

(Table 4: AOR = 0.88). Respondents had lower odds of reporting pain interference in usual activities in 2018 (17%) than in 2014 (21%; AOR = 0.85).

In stratified analyses, reserve soldiers reported a smaller decrease in MSK-MNR (Table C: 0% change; OR = 1.01) and pain interference (1% decrease; OR = 0.97) compared with the full sample. Soldiers with 0 to 5 years of service reported a greater decrease in MSK-MNR (Table C: 6% change; OR = 0.75) and pain interference (5%; OR = 0.72), while soldiers with 6 to 10 years reported a smaller decrease in MSK-MNR (1% change; OR = 1.03) and pain interference (0% change; OR = 0.99) over time compared with the full sample.

DISCUSSION

From 2014 to 2018, we observed decreasing prevalence in some risky health behaviors (i.e., hazardous drinking, tobacco use, weight-loss and performance-enhancing dietary supplements) and medical outcomes (MSK injury, pain interference), along with relatively stable levels of healthy behaviors (dietary habits, regular physical activity, consumption of breakfast and postexercise recovery snacks) within the Army. Sleep quality and duration appeared to improve from 2014 to 2018, although more than one third of respondents reported poor sleep quality or an average duration of 6 hours or less. Differences were also observed among some demographic groups. Women reported better dietary habits and lower levels of risky health behaviors relative to the full sample. Men reported higher levels of physical activity relative to the full sample. Some of these trends (e.g., medical outcomes, alcohol and tobacco use) appeared to

be driven by soldiers who were earlier in their career and were more subtle among those with longer service durations. Soldiers who are early in their career may be more subject to structured routines and healthy lifestyle campaigns aimed at new enlistees than more-experienced soldiers.

Important similarities and differences in health behavior trends exist between civilian and Army populations. Dietary habits remained stable from 2014 to 2018, a finding similar to the general population in which dietary trends have been stable for a few decades²⁰ despite numerous well-publicized campaigns. Many soldiers, particularly junior enlisted personnel, live and work in controlled environments, where environmental and occupational interventions intended to improve health behaviors continue to be tested.²¹ Their diets remain poor overall, which, coupled with short sleep, can compromise readiness and deployability.

Dietary supplement use, including weight-loss and performance-enhancing dietary supplements, decreased over this time period. This trend is similar to those seen in previous surveys, namely the Health-Related Behavior Survey (HRBS)^{7,22} and is encouraging because some dietary supplements marketed for weight loss or body building have been identified in postmarket US Food and Drug Administration warnings as containing potentially dangerous substances²³ that pose a risk to both the general population²⁴ and service members.²⁵ Furthermore, trends in dietary supplement use had previously been higher for military populations compared with civilians, and may be attributable to service members seeking a performance boost to assist with increased occupational demands.²⁶

Results from the National Sleep Foundation's Sleep Health Index questionnaire, administered quarterly to nationally representative samples, showed that 32% of the US population rated their sleep quality as "excellent" or "very good" and 34% rated their sleep quality as "fair" or "poor."²⁷ Since 2016, these levels have remained fairly consistent. The Centers for Disease Control and Prevention reported that approximately 35% of all adults had a short sleep duration on average (less than 7 hours of sleep), with males (35.4%) and adults aged 35 to 44 years (38.3%) having the highest prevalence in 2014.²⁸ Similar to civilians, male soldiers reported slightly lower levels of short sleep duration than female soldiers, though female soldiers reported higher levels of poor sleep quality than male soldiers. Overall, a greater proportion of the Army population reported short sleep duration and poor sleep quality compared with their civilian counterparts,²⁸ although demographic differences complicate these comparisons.

The National Institute on Alcohol Abuse and Alcoholism's National Epidemiologic Survey on Alcohol and Related Conditions shows an increase in high-risk drinking (defined as 4 or more drinks for women and 5 or more drinks for men in a day at least once monthly during the previous year) from 2001–2002 (9.7%) to 2012–2013 (12.6%).²⁹ Historically in the United States, men have consumed alcohol more frequently and in greater quantity than women.³⁰ However, the National Survey on Drug Use and Health, which was completed between 2002 and 2012, reported a narrowing of the gap, including increases in the prevalence of any alcohol consumption in the previous month, the number of drinking

days, and binge drinking among women.³⁰ A greater decrease in hazardous drinking was observed in men than women, which indicates a similar convergence in drinking habits based upon gender.³⁰ These decreases in alcohol use are parallel to decreases observed over the 2011 and 2015⁷ waves of the HRBS, which indicate that our findings may be a continuation of earlier large-scale declines in hazardous drinking.

Our general findings that prevalence of risky health behaviors and self-reported medical outcomes were lower in 2018 than 2014 are important because of their potential impact on Army readiness and deployability. Reductions in risky health behaviors, such as hazardous drinking, may improve overall Army readiness by reducing the number of soldiers deemed nondeployable or being discharged from service because of disciplinary actions.¹⁰ Reductions in self-reported MSK injuries and pain interference may signify some improvements in Army readiness; MSK injuries accounted for 65% of medically nondeployable combat arms soldiers and 70% of medical discharges from 2011 through 2016.³¹ Differences between demographic groups in how the prevalence of health behaviors changed from 2014 to 2018 indicate that changes over time are complex. Future efforts should consider demographic-specific drivers and barriers to positive health-promoting behaviors.

Limitations

All of the measures were self-reported, and, therefore, respondents may be biased toward representing themselves in a positive manner. Even though the data are collected confidentially and

not made available to a soldiers' chain of command, there could still be some inherent distrust about effects their responses would have on one's career.

Also, we cannot directly compare respondents who did and did not allow their records to be used for research. The sample, although large, still represents a convenience sample, unlike the HRBS,^{7,22} which is a sample-stratified survey of active-duty personnel across all branches. Therefore, it is likely that the point estimates from our data may be less accurate than those from HRBS, although, compared with the HRBS, GAT data likely depict a superior view of annual trends and their directions because it was from 1 continuous behavioral surveillance system with minimal changes (at least, with regard to "sampling") over the time span.

Lastly, sample sizes for the 2 models (time-only and covariate-adjusted) differed substantially because 37.9% of respondents were missing administrative data. Respondents with missing administrative data were nearly identical along health behaviors or indicators (percentage differences were within 2%–3%), although they were different along career-related factors: compared with respondents with complete data, those with missing data had more respondents who served for longer than 11 years (21% vs 8%) and fewer who were active duty (42% vs 61%). Nonetheless, only trivial differences were noted in the coefficients for the time-only and fully adjusted models, which indicates that sample sets and person covariates had minimal impacts on estimated changes (i.e., OR associated with year).

Public Health Implications

Health behaviors are a major contributing factor to the health, well-being, and

readiness of the US military. Currently, the US military employs various educational programs and policies to address health behaviors, including (but not limited to) the Army Substance Abuse Program to educate and provide resources for treatment of drug and alcohol abuse, Operation Supplement Safety to provide information on the use of dietary supplements, and the Performance Triad Soldier Challenge, which promotes the importance of sleep, activity, and nutrition. The results of the present study show an overall decline in reported levels of risky health behaviors and steady reported rates of health-promoting behaviors. Despite declines in risky health behaviors, Soldiers still report higher levels of some of these behaviors, such as hazardous drinking and poor sleep quality, than the civilian population. These differences highlight the importance of assessing military-specific health needs. Future research efforts may focus on understanding the direct associations between health behaviors and measures of health and well-being (e.g., physical fitness scores, body composition standards, limited duty rates) as well as the effectiveness of ongoing educational programs and interventions to improve overall health status. In addition, future efforts should assess the impact of new policies and programs (e.g., Army Combat Fitness Test, Army Holistic Health and Fitness program) on improving the health, well-being, and readiness of the US military.

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ABOUT THE AUTHORS

All authors are with the Consortium for Health and Military Performance, Department of Military and Emergency Medicine, F. Edward Hébert School of Medicine, Uniformed Services University, Bethesda, MD. Jessica L. Kegel, Josh B. Kazman, and Daniel R. Clifton are also affiliated with

the Henry M. Jackson Foundation for the Advancement of Military Medicine, Bethesda.

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CORRESPONDENCE

Correspondence should be sent to Josh B. Kazman, 6720B Rockledge Drive, Suite 669, Bethesda, MD 20817 (e-mail: josh.kazman.ctr@usuhs.edu). Reprints can be ordered at <http://www.ajph.org> by clicking the "Reprints" link.

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J. L. Kegel executed original analyses of the data and drafted the article. J. B. Kazman executed additional data analyses and assisted with the write-up. D. R. Clifton, P. A. Deuster, and S. J. de la Motte provided iterative input into the analytic approach and article revisions.

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CONFLICTS OF INTEREST

The authors have no financial interests or relationships to disclose.

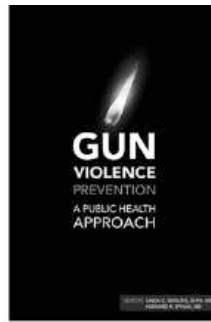
HUMAN PARTICIPANT PROTECTION

The project was reviewed and classified as "research - not human subjects" by the Uniformed Services University Human Research Protection Program Office.


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