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Racial Disparities in Cancer Care | Access to Care for Transgender Veterans | Hepatitis C Among US Veterans | Battling Tobacco Use at Home | HEALTH EQUITY | How Might the Provision of Culturally Competent Services Be Enhanced for American Indian and Alaska Native Veterans? | Improving Trends in Gender Disparities | Suicidality Among Hispanic and African American Veterans Following Surgery

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COVER: Two-year-old Alejandro sits in his stroller while his mother (rear), who asked not to be identified because of her immigration status, fills out paperwork for help with her rent payment at La Colaborativa in Chelsea, Massachusetts, March 23, 2021. La Colaborativa has adapted to help the residents of Chelsea, one of the US cities hardest hit by the COVID-19 pandemic, by operating a food pantry, providing a vaccine clinic, and offering housing assistance as people face evictions, which have continued despite moratoriums.

Cover concept and selection by Aleisha Kropf. Photo by REUTERS/Brian Snyder. Printed with permission.

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Keeping Public Health Advocacy Strong

I am often asked what keeps me up at night. My responses are varied, of course, but they include a myriad of issues. We have lost so much ground this year on women's reproductive rights. My heart aches for women and their families during this time of difficulty. Our rural communities are still hurting and have had little to no authentic attention to their needs, despite what the spotlights of the COVID-19 pandemic and the opioid crisis have shown. Rural health is a major public health issue that requires concerted coordination between the federal, state, and local governments as well as the public in general to address.

We simply cannot keep assuming it is "just the way it is." We have begun to make progress on climate change but only after years of not being able to have those discussions openly, much less adequate funding to address it. Our public health workforce is eager to work on this issue in whatever ways best resonate with their communities. The increase in gun violence in our country is a symptom of deeper issues that desperately need attention so that we do not keep losing young members of our communities.

Being elected as the president of the American Public Health Association (APHA) is one of the most awesome gifts that one can be given. Serving in this capacity in 2022, I was able to travel to several state public health association meetings and to interact directly with colleagues from all over the country. I was committed to encouraging the public health workforce after their long and seemingly never-ending challenges of the pandemic. I found them to be exhausted mentally and physically from the pandemic's demands but also dedicated to public health going forward. Although I recognize that we have lost vital members of our workforce during this time, we still have many who are excited about the public health mission in their jurisdictions and are gearing up for the next challenges. I am heartened by that enthusiasm!

I support the national reports that recommend changes to our public health system and support to our public health workforce going forward (https://bit.ly/3LNha0a; https://bit.ly/3CafqtR; https://bit.ly/3SI2kut). May these reports lead to action that positively affects our state and local public health systems, and may those changes begin to happen soon.

I am also heartened by the students and graduates who have entered the profession with a new way of thinking and a commitment to making public health better and stronger. Our state affiliates have strong student membership numbers, or they are working on achieving them. That strengthens my faith in those who will be our leaders in the future!

Finally, if we really are committed to the social determinants of health, we must address the issues that affect public health locally. Clean water, accessible health care, accessible public education, good roads and public transportation, and political will to consider the health of all of the people living in our country are basics. We cannot take our eyes off the ball. The APHA's incoming president, Chris Chanyasulkit, PhD, MPH, often says that voting matters, and she is right. As APHA members and public health professionals, we have to consider all of the factors that affect public health and not lose our strong advocacy for those issues. We still have much work to do. AJPH

Kaye Bender, PhD, RN
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American Public Health Association President, October 2021–November 2022

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3 Years Ago

Ensuring Compliance With Quarantine by Undocumented Immigrants and Other Vulnerable Groups: Public Health Versus Politics

[M]any unauthorized parents and children lack trust in the immigration system. They may experience such stress or fear about separation from their families that they even decline to receive necessary medical care. Many fear that if they cooperate with any emergency measures public health officials will learn their citizenship status and report them to local police or ICE. Another complicating factor is that undocumented immigrants are excluded from public insurance programs, such as Medicaid, as well as subsidies under the Patient Protection and Affordable Care Act. They may avoid applying for private insurance coverage because of the "perceived or actual need to show documentation of immigration status." . . . These barriers to obtaining affordable health coverage exacerbate existing health disparities in this vulnerable population.

From AJPH, September 2019, p. 1180

14 Years Ago

Immigrant Children's Reliance on Public Health Insurance in the Wake of Immigration Reform

Contrary to popular perceptions, foreign-born children in the United States do not rely on public health insurance programs more than US-born children, despite reversal of the public charge rule. Even after the significant socioeconomic differences between US-born and foreign-born children had been taken into account, the vast majority of foreign-born children in our study were much more likely than were US-born children to be uninsured, to be living in poverty, and to have parents with less than a high school education. Such cumulative social disadvantage is likely to adversely affect the ability of immigrant children to become productive members of the American labor force. In the various discussions of proposals for universal child health coverage, policies designed to promote the healthy growth and development of this highly underserved population merit serious consideration.

From AJPH, November 2008, p. 2007

Using Low-Cost Sensor Networks: Considerations to Help Reveal Neighborhood-Level Exposure Disparities

Angie Shatas, MS, and Bryan Hubbell, PhD

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See also Esie et al., p. 1765.

The growing availability of low-cost sensors can potentially democratize the process for reducing disparities in exposures to harmful air pollution. When used collaboratively with government agencies and researchers, sensors deployed by community organizations can build trust in environmental decision-making.[1] Low-cost continuous sensors can complement regulatory monitoring networks required by the Clean Air Act, which have high confidence but relatively low geographic coverage. Sensors are often portable or even mobile and can prove particularly useful if they measure some of the same air pollutants as regulatory monitors.

Sensors deployed on a neighborhood scale can reveal spatial and temporal variations in air quality, as Esie et al. (p. 1765) show. Increased temporal resolution can identify episodes of poor air quality that exacerbate existing inequities in exposure or if those episodes, when compared with mean air quality levels, create new inequities. Increased temporal resolution can show when exacerbations happen and, in combination with higher spatial resolution, can then reveal the cause. Identifying the sources of the emissions provides communities and decision-makers with the information needed for action in addressing inequities. However, communities and government agencies must work together to agree on how to interpret and evaluate sensor data, especially in cases when it may not agree with regulatory monitors, to prevent friction and loss of trust.[1]

Sensors have a wide appeal and market availability, but the quality of data they generate must be considered. To help those using sensors as part of air monitoring, the US Environmental Protection Agency's (EPA's) Air Sensor Toolbox[2] provides the latest science on sensor performance and operation, and the EPA is providing $20 million in grants to enhance community and local efforts in monitoring air quality, including in or near underserved communities.[3] In addition, the Inflation Reduction Act[4] contains provisions to deploy air monitoring in communities, including deploying sensors in low-income and disadvantaged communities.

SELECTION OF DATES, TIMES, AND LOCATIONS IS CRITICAL

Even for low-cost sensors, air quality measurement campaigns can be resource intensive, and thus decisions often need to be made about where and when to take measurements. Esie et al. conducted measurements for July 2021 because July had historically shown higher fine particulate matter (diameter $\leq$ 2.5 µm; $PM_{2.5}$) levels. Although overall $PM_{2.5}$ levels were below daily standards, there were relatively elevated $PM_{2.5}$ measurements, predictably on July 4 and unexpectedly on July 23 because of a wildfire smoke incursion event. The latter event showed minimal variation across neighborhoods with different sociodemographic profiles. Summer months often show a high contribution from regional sulfate from power generation (although this contribution has fallen over time), and, as such, more local contributions may be masked. Looking at other months may have revealed more significant disparities across neighborhoods, perhaps because of greater proportional contributions from local industries or from urban transportation or differences in heating emissions. Recent trends show that in many regions of the country, including Chicago, relative peaks in $PM_{2.5}$ now occur in the winter, and those peaks may be associated with more local emission sources.[5] Other temporal events of concern would be short-term sources of emissions from industrial sources (such as shutdown/startup malfunctions or maintenance), particularly if those sources are proximate

AJPH December 2022, Vol 112, No. 12

to communities with environmental justice concerns. Temporary increases in emissions such as these would be both isolated in time and space, in contrast to the two events in July.

For the purposes of understanding exposure disparities, focusing on spatial or temporal excursions from mean total $PM_{2.5}$ levels may be more useful than looking at total $PM_{2.5}$. Identifying a local "hot spot" that might contribute to disparities in exposure would entail subtracting citywide, regional, and national contributions until only the excess $PM_{2.5}$ associated with local contributors remained. An approach has recently been proposed to remove regional background and provide a decomposition of $PM_{2.5}$ air pollution into long-range, midrange, neighborhood, and near-source for all census tracts in the United States,[6] and this approach may also remove autocorrelation in a more structural way rather than using spatial lags. Further removing longer-term temporal trends from these spatially decomposed $PM_{2.5}$ levels would highlight temporal excursions that may also lead to additional disparities. In both cases, the analyses not only would identify when and where disparities occur but also could help to diagnose the emission sources that cause the disparities.

Siting of a network of low-cost sensors can be focused on diagnosing where and when inequities in exposure occur and on identifying the cause(s) of the inequities. The siting should be done with community input. Esie et al. used sensors located at bus stations, which are convenient locations and could capture near-road $PM_{2.5}$ exposures. However, these locations might not be best for identifying $PM_{2.5}$ exposures from industrial or other sources.

UNDERSTANDING DISPARITIES REQUIRES EQUITABLE NETWORKS

When properly sited, and with a dense-enough sensor network, it becomes possible to predict $PM_{2.5}$ levels at other neighborhood locations. For example, a community may wish to identify places of neighborhood concern or places with sensitive populations. Inverse distance weighting (IDW) or cokriging approaches that incorporate additional information such as wind directions[7] can provide spatially resolved predictions that are similar in quality to land use regressions or downscaled model predictions.[8] However, it is not clear that the density or location of sensors at bus stops satisfies criteria for using IDW as employed by Esie et al., and thus statistical models that seek to identify the disparities in $PM_{2.5}$ concentrations across different races using IDW may suffer from exposure misclassification. To respond to the concern of Esie et al. about temporally invariant covariates, it may be possible to use new land use regression methods[9] that allow both spatial and temporal decomposition.

Esie et al. importantly note that crowdsourced sensor networks tend to be located in White, high socioeconomic status neighborhoods. If higher-income neighborhoods have more access to air quality sensors and more ability to respond to the information they generate, disparities in air pollution health outcomes can be exacerbated.[1] This reveals a need for more consistent, government-sponsored networks, which could promote interoperability and equitable access. By allowing a cross-comparison of data gathered using disparate sensor networks, information could be compared and shared on a broader scale.

CONCLUSION

The Esie et al. study adds to evidence that disparities in exposure continue to exist in Chicago and that certain types of emission events can exacerbate those disparities. The types of emission events identified are difficult to regulate, and the study design is not able to identify harder-to-diagnose sources of air pollution excursions. A greater focus on the times and places that have substantially higher neighborhood air pollution levels would advance two goals: a greater ability to ascertain the sources of inequities and information that can empower communities working with government agencies to prevent those emission events and reduce exposures. Finally, low-cost sensors, with their affordability and ease of deployment, have the potential to collect data that can reveal air quality and exposure disparities, but the data will have the most impact in rectifying disparities when communities and government agencies agree, preferably in advance, on how to evaluate and interpret the data. AJPH

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The authors have no conflicts of interest to declare.

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Effect of Masking to Prevent COVID-19 Transmission in Schools and the Responsibility of States to Protect Public Health

Chloe A. Teasdale, PhD, and Sasha A. Fleary, PhD

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See also DeJonge et al., p. 1791.

The 2022–2023 school year marks the third time US children and adolescents have returned to school during the COVID-19 pandemic, and there is hope that it will be less challenging than the previous two years. At the start of the 2021–2022 school year, the highly transmissible Delta variant was causing rapid increases in cases and hospitalizations, notably among children and adolescents.[1] By December 2021, the even more infectious Omicron variant had emerged. At its peak in January 2022, Omicron caused almost 1 million cases per day, with a rate of new cases among school-aged children (5–11 years) of 1545 per 100 000 per week.[2] The Omicron surge also caused considerable disruption of school attendance. In a *New York Times* poll, half of US parents reported that their child missed three or more days of in-person schooling in January 2022.[3]

The COVID-19 pandemic may not be over, and there is considerable uncertainty about how COVID-19 will impact the current school year and about our level of preparedness. Given the pattern of the past two years, the potential for new surges is an important concern, particularly given the low vaccination coverage among school-aged children. Although COVID-19 vaccines are now authorized for children aged six months and older, at the start of the school year, only 31% of school-aged children had been vaccinated, and less than 15% had received a booster.[2] Uptake of boosters among adults is also low, even among those at high risk because of age or comorbidities.

An additional concern is that the guidance on COVID-19 mitigation strategies from the Centers for Disease Control and Prevention (CDC) for the 2022–2023 school year recommends a localized approach that is responsive to COVID-19 community levels as indicated by hospitalization and case data. The CDC recommends using face masks in schools when community infection levels are high[4]; however, our recent work suggests that implementation of these guidelines may be challenging. Because of the widespread use of at-home antigen tests, which are not included in routine surveillance data, it may be difficult to accurately track the number of new COVID-19 cases and to use these data to rapidly respond to surges.[5]

EFFECTIVENESS OF SCHOOL-BASED COVID-19 PREVENTION STRATEGIES

In this issue of *AJPH*, DeJonge et al. (p. 1791) present the findings of a study comparing the effectiveness of COVID-19 mitigation strategies in Wisconsin school districts in the fall of 2021. Using employment records and COVID-19 testing data from September through November 2021, they compared the incidence of new cases of COVID-19 among teachers working in school districts with prevention policies with the incidence of infections in districts without policies. The researchers examined the individual effects of three different COVID-19 prevention strategies: mask wearing by teachers and students, physical distancing, and quarantine.

The study found that the overall COVID-19 incidence rate was 5458 per 100 000 educators during the first three months of the 2021–2022 school year. The researchers also showed that although distancing and quarantine had no impact on reducing infections among teachers, masking policies were associated with decreased risk of infection. Teachers across all grade levels who worked in districts with masking policies were 19% less likely to have a positive test result for COVID-19 than those in districts without masking

AJPH December 2022, Vol 112, No. 12

policies (hazard ratio = 0.81; 95% confidence interval = 0.72, 0.92). Furthermore, the study shows that even among a highly vaccinated population (78% of Wisconsin teachers were fully vaccinated), masks were protective against COVID-19 transmission. These findings demonstrate that during a period of high infection rates, the combination of masking and vaccination provided stronger protection than vaccination alone.

The study is comprehensive and has important strengths. It includes data from 307 Wisconsin school districts (81%) and almost 52 000 teachers. It adjusted for critical confounders, including the age, sex, and vaccination status of teachers, as well as community characteristics (vaccination coverage and infection rates) and school-level factors (average class size and location). Notably, it was conducted before the widespread use of at-home antigen testing, which could make conducting similar studies more difficult because of decreased recording of cases.[5] A reported limitation of the study is the lack of accounting for adherence to COVID-19 prevention policies by districts. However, nonadherence to prevention policies would most likely have biased the results toward the null, indicating that masking in schools may be more protective than this study was able to show.

Unfortunately, this study did not measure COVID-19 infections among students to demonstrate the direct benefit of mitigation strategies for children and adolescents. The finding of reduced infections among educators is indicative of lower transmission within schools, which is indirect evidence of the impact on students. The findings are consistent with previous studies showing that masking prevents secondary transmission in schools.[6] This evidence for the protective effect of masking in the school environment is important and timely, given the high levels of COVID-19 vaccine hesitancy among US parents that we have previously reported.[7,8] With so few school-aged children vaccinated, these findings are particularly relevant because masking will be a critical prevention intervention in the event of another COVID-19 surge during the 2022–2023 school year.

RESPONSIBILITY OF STATE GOVERNMENTS TO PROTECT PUBLIC HEALTH

The study by DeJonge et al. also demonstrates how many US states refused to implement evidence-based public health policies that would have protected their workers, students, and communities during a critical point in the pandemic. At the start of the 2021–2022 school year, only 18 states had mandates requiring masking in schools, whereas eight states passed laws prohibiting school districts from requiring masks, and the remaining 24 states allowed local decisions about masking policies.[9] KFF reported that in the fall of 2021, more than two thirds of school-aged children lived in US states that either did not have mask requirements or prohibited them. Explanations for why several states chose to legislate against evidence-based COVID-19 mitigation strategies have been examined in previous editions of *AJPH*.[10] Fewer studies have described the reasoning behind and consequences of the approach taken by other states that left critically important public health decisions up to individual school districts.

Wisconsin was one of the states that did not adopt a statewide mask mandate for schools in the fall of 2021, in spite of CDC guidance recommending the use of masks and existing evidence at the time showing lower COVID-19 caseloads in states that implemented mask mandates.[11] DeJonge et al. showed that at the start of the 2021–2022 school year, the most common COVID-19 mitigation policy in place across Wisconsin school districts was physical distancing, adopted by 68%, followed by quarantine, implemented in just over half (52%). Only 25% of school districts in Wisconsin in the fall of 2021 had masking policies, while 21% of school districts were not implementing any of the COVID-19 mitigation policies examined.

According to the Wisconsin Department of Health Services, COVID-19 cases were increasing in early September 2021 and continued rising throughout the fall.[12] The largest increase in cases during this period was among school-aged children and adolescents, which marked the first time that COVID-19 cases in children across the state had outpaced those in adults.[12] In the week of September 12, 2021, those younger than 18 years in Wisconsin had a COVID-19 infection rate of 447 cases per 100 000 (5624 cases) compared with the next highest age group, 35–44-year-olds, with 345 cases per 100 000 (2492 cases). Wastewater surveillance from this time showed similar trends of rising infections across the state after the start of the school year. In this context, it is remarkable that more was not done at the state level to protect Wisconsin's students and educators from COVID-19. Further studies should explore whether lack of statewide mandates compounded unequal distributions of COVID-19 cases, hospitalizations, and deaths.

The study by DeJonge et al. adds to the body of evidence showing that the use of face masks helps prevent COVID-19 transmission in schools and communities.[11,13,14] These data are critical for informing plans for future surges, when widespread use of masks may be necessary again to protect children, educators, and their communities. In addition, enhanced surveillance that does not rely solely on reported cases is also needed to allow immediate and appropriate interventions. Finally, this study demonstrates that reliance on local decision-making about critical public health measures left many schools unprotected from COVID-19 and created inequities in risk for Wisconsin's children and educators. Protecting public health is one of the fundamental responsibilities of governments, and the COVID-19 pandemic has made it clear that many state governments need to take stronger actions to protect the health of all of their citizens. AJPH

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Living Alone and Suicide Risk: A Complex Problem Requiring a Whole Population Approach

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See also Nestadt, p. 1702 and Olfson et al., p. 1774.

Death by suicide is one of the great challenges in public health. Suicide is a tragedy that affects not only the deceased individual but also everybody to whom that individual was connected. Yet, despite the link between death by suicide and social integration being long recognized and many efforts to reduce suicide rates in recent decades,[1] the age-adjusted rate of suicide in the United States increased from 10.5 per 100 000 in 1999 to 13.9 in 2019.[2] Before the pandemic, suicide was the 10th leading cause of death in terms of all-age mortality; in comparison, age-adjusted rates for the top three causes of death are 161.5 per 100 000 for heart disease, 146.2 for cancer, and 49.3 for unintentional injuries. Suicide is even more important among those younger than 65 years, ranking as the fifth leading cause of death.[2] The US suicide rate is not atypical, with the suicide rate for other high-income countries being 13.7 per 100 000.[3] Thus, identifying how to best target interventions to address suicide is a global priority.

In this issue of *AJPH*, Olfson et al. (p. 1774) describe who dies alone and how. The strongest associations between living alone and risk of suicide are for those with the most advantaged social positions, as indicated by education, income, and ethnicity. Looking at the results in additional detail provides more information. Among adults living with others, suicide rates decline with increasing income and education levels. Conversely, there is little evidence of any differences in suicide rates by income or education among people living alone, a finding that cannot be explained by chance.

Essentially, living alone, particularly in the case of men, seems to be associated with not only an increase in the risk of death by suicide but also an absence of a social gradient in death by suicide. Given the acknowledged lack of adjustment in the Olfson et al. study, it is possible that living alone could be a marker for previous mental health issues or other factors. However, the relationship between death by suicide and living alone has been shown to persist after adjustment for poor mental health[4] and merits further discussion.

Two possible theories stand out in explaining the Olfson et al. results: the concept of "thwarted belongingness" from the interpersonal theory of suicide and the integrated motivational–volitional theory. Thwarted belongingness is the perception that a person is alone and lacking any reciprocal caring relationships, and clearly living alone is potentially a marker for thwarted belongingness.[5] However, according to the overall theory, thwarted belongingness alone is not sufficient to induce suicidal behavior. The Olfson et al. study lacks indicators for another necessary component of the theory, "perceived burdensomeness," which indicates the degree to which people feel liability to others or self-hatred, and so other theories are required.

The integrated motivational–volitional theory divides the development of suicide into three different phases: the premotivational phase, the motivational phase, and the volitional phase.[6] In the premotivational phase, background factors set up vulnerabilities to suicidal behavior; these factors include negative life events, social circumstances, and biological factors that might predispose people to suicidal behavior. One characteristic in the premotivational phase that has been consistently linked to suicidal behavior is socially prescribed perfectionism, defined as people's belief that others hold unrealistically high expectations of them.[7] Perfectionism is certainly something that could drive people to be educationally successful and pursue higher incomes and that could increase the risk of suicide. This alone would not be enough to explain the Olfson et al. results.

The integrated motivational–volitional theory also puts forward that those predisposed to suicide do not automatically progress to suicidal ideation or intent. In the motivational phase, it is argued that events or situations may arise that induce feelings of defeat or humiliation, ultimately leading people

AJPH December 2022, Vol 112, No. 12

to a feeling of being trapped with no perceived escape.[6] Factors such as a failure to achieve goals, thwarted belongingness, a lack of coping skills, and lack of availability of social support are proposed to affect transitions to suicidal intentions. In this context, living alone may not simply be an indicator of loneliness or thwarted belongingness but also an indicator of failing to achieve important life goals and the support thereby obtained, such as having a partner or raising children.[4]

The volitional phase is when thoughts about suicidal intent turn into actions. Factors that are important at this stage include access to means to complete suicide, exposure to and knowledge about suicide, and personality traits such as impulsivity.[6] As discussed by Olfson et al., living alone may be more strongly related to suicide by poisoning because it limits the opportunities for other people to intervene. Although the Olfson et al. results are consistent with theory, the lack of data on important factors such as mental health indicate that there are alternative explanations. Their study provides further support for the integrated motivational–volitional theory, but the entire process has not been tested.[6]

Developing interventions that address issues such as perfectionism and failing to achieve family goals will not be easy, and there might be tradeoffs in terms of economic outcomes. Thus, addressing suicidal behavior requires a recognition that it is a complex issue necessitating different methods and study designs. Methods may include ecological momentary assessments, which enable data to be collected in real time and may provide better information on specific components of suicidal behavior.[6] In addition, social network analysis may provide insights into social connections and how living alone relates to suicidal behavior.[6]

New approaches such as simulations may be required to synthesize data from multiple sources.[8] In this context, descriptive studies such as that of Olfson et al. can be used to validate the results of simulations informed by other designs. However, given the difficulty of predicting suicidal behavior[6] and the commonness of important risk factors such as living alone, targeting high-risk groups may not be practical without further research.

An alternative to targeting interventions toward high-risk individuals would be to take a whole population approach,[1] as suggested by Geoffrey Rose.[9] Restricting access to the methods used for suicide appears to be successful in reducing suicide rates, as there seem to be limited substitution effects.[10] The Olfson et al. results confirm a suitable target for intervention. Firearm deaths contribute to more than half of suicides irrespective of living arrangements. Although it is not easy to generate the political will to limit access to firearms, policies such as mandatory waiting periods and background checks have been shown to reduce suicide rates.[11] Other possible options include limiting the number of tablets included in a packet of paracetamol and restricting access to suicide hotspots such as bridges.[10]

However, the appropriateness of interventions is context dependent. In addition, caution is needed in interpreting the results of interventions; for example, the implementation of restrictions on the package size of paracetamol tablets took place in combination with other polices aimed at improving health more generally. Consequently, the relationship between reduced package sizes and reduced suicide rates may be confounded by other policy changes.[12] Reducing suicide rates may require approaches aimed at improving the whole population's health in general rather than simply decreasing suicides. AJPH

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Suicide and the Solitary Life: Differential Risks of Living Alone Across Sociodemographic Groups

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See also Shaw, p. 1699 and Olfson et al., p. 1774.

At 77 years old, even after decades of prodigious philanthropy, George Eastman remained one of the wealthiest men in the world. The unmarried founder of Eastman–Kodak lived alone until March 14, 1932, when he revised his will in the presence of his lawyers, dismissed them from his study, folded a wet towel over his chest, and shot himself through the heart with his desk drawer revolver.[1] His obituary reported, "A sense of loneliness encompassed George Eastman, after the recent deaths of two of his closest friends, and led him to take his own life."[2(p5)]

Living alone, loneliness, and social disconnection have been proposed as suicide risk factors since the dawn of suicidology.[3] However, a lack of predeath data on large samples of suicide decedents has prevented us from knowing the demographic characteristics of those at highest increased risk when living alone. A new study by Olfson et al. in this issue of *AJPH* (p. 1774) contributes evidence of the association between living alone and suicide as it varies across demographic and socioeconomic subgroups. The authors reviewed the 2008 American Community Survey, which includes more than 3 million adults linked to the National Death Index, to identify suicide deaths over the 11 succeeding years. The participants reported on their living situation as well as sociodemographic characteristics, self-reported disability, and housing information, including residential stability and homeownership.

Olfson et al. found the annual suicide rates of adults living alone to be almost twice that of adults living with others, confirming previous reports.[4,5] The authors went on to identify large differences in the strength of that association across specific subgroups. The associations between living alone and subsequent suicide were found to be strongest among wealthy, well-educated, male, White, and older age groups. Membership in some of these groups was previously known to independently increase suicide risk,[6] and their strong associations with living alone is tragically reminiscent of George Eastman. However, the recognition of low social integration as a risk factor for suicide dates back most prominently to Emile Durkheim's investigations in the 19th century.

In his landmark book *Suicide*, Durkheim cited the 1886 French census in pointing out that the lower the average number of persons living in the family home, the higher a region's suicide rate.[3] He raised this as a central tentpole of his theory of "egoistic" suicide, which is undertaken by those who see themselves as alone or disconnected from socially integrated groups. Egoistic suicide is thought to be more common in less socially integrated communities but is also noted to be associated with certain types of individuals in a given society. For instance, Durkheim posited that being unmarried or widowed was associated with increased suicide risk. This went against the earlier belief that marriage was the higher risk state, a finding that resulted from past failure to adjust for age in comparing married to unmarried individuals.

Like Olfson et al., Durkheim also related suicides of social isolation to the attainment of knowledge and education, although he did so indirectly by pointing to differential levels of education in distinct religious groups and their associated suicide rates at the time. He credited the higher rates of suicide among Protestants to their greater "pursuit of free inquiry" and learning compared with Catholics, who had a much lower suicide rate. Durkheim argued that this free inquiry steered some Protestants further from their church communities, resulting in weakened community bonds and more vulnerability to suicide. He further performed some intellectual gymnastics to explain the lower rates of Jewish suicides, despite higher levels of education, as evidence that Jewish education is in line with their religious doctrine and so serves to further socioreligious integration. However, in view of our modern understanding of stigma, it may be more likely that the stronger condemnation of suicide by Jewish and Catholic leaders provides a better explanation for the lower suicide rates in those groups.

Aside from education level, Olfson et al. found the strongest association between living alone and suicide existed in high earners. In general, suicide risk is greater in persons experiencing poverty or homelessness.[7] However, in the context of living alone, Durkheim suggested an explanation for increased suicide among the wealthy. He theorized that the wealthy depended less on others for material support and, thus, felt less invested in the larger community. Durkheim wrote that for most, interdependency in a group creates a reciprocal investment in others that prevents one from being overwhelmed by one's own troubles and contextualizes them in larger communal joys, hopes, and a future. This allows a suffering individual to "share in collective energy and support his own when exhausted." By contrast, the wealthy individual may feel they owe society nothing and "have no reason to endure life's sufferings patiently."[3(p168)]

The recognition of social integration as suicide prevention did not end with Durkheim. Thomas Joiner's interpersonal theory of suicide[6] incorporated the concept of "thwarted belongingness" in recognition of the increased risk of an unmet need to belong. Thwarted belongingness is thought to partially explain the association between suicide and living alone[8] as well as its associated corollary, loneliness.[4] Rory C. O'Connor's integrated motivational–volitional model of suicide continued to develop this idea by highlighting loneliness as a key moderator between a sense of entrapment and subsequent suicidal acts.[9] These theories persistently recognize the importance of social integration because being alone continues to be identified as a risk factor for suicide both directly and as a contributor to mood disorders.[4]

Although the psychological impact of living alone and loneliness may add to suicide risk, there are also practical considerations to account for when considering the risks of living alone. In a secondary analysis, Olfson et al. found that the association between living alone and suicide varied significantly by suicide method. Poisoning, which accounts for most suicide attempts in the United States but a minority of suicide deaths,[10] demonstrated the strongest association. In comparison, for firearm suicides (the most common method of US suicide), living alone was less strongly related to suicide risk. This may be unsurprising, given that suicide attempts by poisoning leave time and opportunity for rescue by a housemate, whereas in firearm suicide attempts, rescue is usually impossible.

Safety planning interventions recognize access to lethal means as a prominent risk and suggest the use of social contacts both for emergency support and for making the environment safer by eliminating access to lethal means.[11] A recent study of veterans found that lack of social contacts on the safety plan was associated with more than double the risk of subsequent suicidal acts, further highlighting the role of social integration in practical safety considerations.[12]

Of note, this study was unable to exclude some important potential confounders of the association between living alone and suicide. Psychiatric illness, a major risk factor that was largely underappreciated by Durkheim, could not be reliably measured in this sample. Mood, anxiety, and substance use disorders have been independently associated with both suicide and living alone,[6,13,14] and so we cannot be certain that there is a causal relationship between living situation and subsequent suicide without these diagnoses included as covariates. However, as the authors point out, several previous studies have found that the association holds even when psychiatric morbidity was included in the models.[4,5]

The findings of Olfson et al. bolster more than a century of work underlining social isolation's association with suicide. By focusing on the objective measure of living alone, as opposed to the more difficult to quantify and evaluate concept of loneliness, the authors present clinicians with a potential risk factor that is easily identified in patients and can be integrated into existing risk stratification strategies. Beyond that, living alone is a modifiable risk factor that can be addressed by public health and social work interventions, much as we can address other major suicide risk factors, such as poverty, psychiatric illness, and lethal means access. AJPH

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Promoting Evidence-Based Policy Solutions to the US Gun Violence Epidemic

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See also Rowhani-Rahbar et al., p. 1783.

Firearm ownership and carriage in the United States remains an intensely personal, political, and social issue. In recent years, firearm ownership has remained at the forefront of our society's dialogue about safety, personal freedom, and the role of government in regulating firearm access. Although some national regulations govern access to firearms, specific restrictions and regulations on firearm ownership and carriage are largely delegated to the states. Because states have varying laws, there are many opportunities for natural experiments that explore the relationship of firearm regulations to firearm ownership and carriage behavior, injury, and mortality.

In public health, we are interested in the external, population-level interventions and policies that decrease death and disability from injury or disease. Oftentimes, behavior change as a result of an intervention or policy is the crucial step that prevents the negative health outcome. Consequently, Rowhani-Rahbar et al. (p. 1783 in this issue of *AJPH*) have sought to understand whether state-level differences in firearm policies affect firearm carriage behavior with a loaded handgun, with an understanding that such behavior may be linked to the outcomes of interest, in this case, injury and death by firearm. It is known that firearm access and carriage are two of the most significant risk factors for pediatric firearm injuries,[1–5] intimate partner homicide,[6–8] suicide,[9–11] and homicide of those who cohabitate,[12–14] although relatively less is known about the relationship of population-level firearm carriage to population-level death and injury by firearm.

By using a nationally representative sample survey of firearm-owning adults, Rowahani-Rahbar et al. analyzed loaded handgun carriage. They then described the groups of respondents by demographics and the reasons cited by the owner for carrying the weapon. They also demonstrated that firearm owners carried loaded weapons in significantly more permitless carry and shall issue states than states with may issue policies. These data, when extrapolated, demonstrate that about 16 million adults in the United States have carried a loaded handgun in the past 30 days, a significant increase over 2015 data, which estimated 9 million adults did so.[15]

This research is critical for state policymakers to review. These data show that specific state-level policies can decrease the carriage of loaded handguns among those state populations. Rowahani-Rahbar et al. have demonstrated that policies aligned with may issue firearm carriage permitting can decrease the number of child, adolescent, and intimate partner homicides; lawmakers who wish to decrease these deaths and whose constituents wish to decrease the population level of loaded gun carriage will be interested in these results. Furthermore, it is critical for those in the public health field to continue to perform high-quality and meaningful research on the implications of public policy on injury and public health. The federal and state governments should continue to provide funding mechanisms for such research to ensure that our policies are evidence driven and scientific in their approach to reducing injury and death. AJPH

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The Perfect Gun Policy Study in a Not So Perfect Storm

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See also Kapadia, p. 1710.

In 2019, Louis Klarevas, Andrew Conner, and David Hemenway published "The Effect of Large-Capacity Magazine Bans on High-Fatality Mass Shootings, 1990–2017."[1] This seminal study empirically demonstrated that prohibition of large-capacity magazines (LCMs) attenuates mass shooting incidents and lethality.[1] The article ranks in the top 1% of high attention scores and is the most cited and discussed research study in social and legacy media in the history of *AJPH*. To date, the study has been mentioned in 569 media sources including 73 news outlets (with 87% of the mentions being made by the general public), and there have been 32 research citations.[2] Dimensions, a research insights platform, reports that the article has received approximately eight times more citations than average.

To that end, we explain why this study continues to have a large impact, leaving an indelible mark in academic circles while garnering the public's attention despite the political, academic, personal, and cultural hurdles Hemenway has faced dating back to the 1990s.

THE POLITICIZATION OF SCIENCE

Although many topics have been politicized, public health research on gun control was intentionally suppressed by the federal government through the Dickey Amendent.[3] The 1996 congressional appropriations bill stipulated that "none of the funds made available for injury prevention and control at the [Centers for Disease Control and Prevention] may be used to advocate or promote gun control."[3(p549)] The political fallout and academic witch hunts, combined with a dearth in funding, were unprecedented.[3–5] A handful of academics, people such as Hemenway, kept the lights on and continued to link policies to gun injury prevention.[3,5] They challenged our thought leaders at a time when there was not enough political will[6] to reduce gun violence. These academics became icons because they did their research in a hostile environment. They lost research funding, were targeted by the National Rifle Association, and faced daunting congressional inquiries.

Violence researchers either had to remove guns from their research or risked being defunded or attacked by the US Congress and gun rights advocates. Thankfully tenure prevailed, or even the scant public health gun studies would never have happened. The historical context of politicization elevates Klarevas et al. because earlier work, especially that of Hemenway, was published under attack, much like the work of scientists who study climate change, critical race theory, or COVID-19 masking. However, Klarevas and Hemenway have published other research on gun violence that did not rise as high on the public agenda, so we must look to additional factors to understand what catapulted this particular study.

PUBLIC HEALTH IMPLICATIONS OF THE SUPREME COURT RULING

On June 23, 2022, the Supreme Court ruled in favor of the Second Amendment's operative clause (the right of people to keep and bear arms shall not be infringed) over the prefatory clause (a well-regulated militia being necessary to the security of a free state). The court expanded individual gun rights and threw out several lower court rulings that upheld gun restrictions, including bans on assault-style rifles in Maryland and large-capacity ammunition magazines in New Jersey and California.[7] In addition, the court limited state policies regarding the purchase, possession, and transportation of firearms and revoked the only gun-control policies known to curtail mass shootings. Thus, the number of mass shootings and their lethality will continue to rise. Ironically, the search for some secret solution to stop mass shootings will redirect policymakers and journalists back to Klarevas et al. yet again after the next mass shooting.

NEWS COVERAGE OF MASS SHOOTINGS

In 2021, nearly 49 000 people in the United States died from guns.[8] Another 100 000 were shot but survived their injuries. Approximately 60% of gun

AJPH December 2022, Vol 112, No. 12

deaths were suicides, less than 5% were accidental or police shootings, and approximately 36% were homicides. Most homicides are not the result of mass shootings.[9] In fact, mass shootings make up less than 1% of gun deaths but account for most of the media attention. Suicides are more commonplace events but less likely to make the national news cycles.

The result is that mass shootings more often appear as news stories, thus distorting the public's perceptions about which types of gun deaths are bigger threats. Parents are more terrified that their children will be killed in a school mass shooting even though there is a much higher likelihood that they will injure or kill themselves if there is an unsecured gun in the home. In the aftermath of a mass shooting, gun violence researchers become part of the media frenzy and are interviewed at length. Because research on mass shootings crosses over from the scientific community to the public, mass shooting research is more likely to capture the general public's attention.

A BUNCH OF FIRSTS AND SCIENTIFIC RIGOR

First Hemenway and later Klarevas published public health gun research when it was unpopular. More important, they set the stage for gun violence researchers who would come after. Klarevas et al. published their article in 2019, just before the 2020 federal budget included $25 million for the Centers for Disease Control and Prevention and the National Institutes of Health for research on reducing gun-related deaths and injuries after a 24-year hiatus, paving the way for a proliferation of new gun violence research.[10,11]

MASS SHOOTING RESEARCHERS DO NOT AGREE ON MUCH

Within academic circles, there is much debate about what constitutes a mass shooting, where it happens, how many people die, and which data to use. Regardless, Klarevas, Conner, and Hemenway followed the public health standard for how to do policy-relevant research. First, they built on existing science on gun violence and mass shootings. Second, they isolated a specific type of mass shooting, one with high lethality (six or more fatalities), and then linked policy to prevention. Gun violence and mass shooting researchers cited this study because the authors used a narrow and specific definition of high lethality, including number of people killed, where the shooting occurred, by whom, the data source, and inclusion and exclusion criteria.[11]

Even the Federal Bureau of Investigation does not have a mass shooting definition. Instead, it defines "mass murder" as an incident in which four or more people are killed, which can include gun violence. Klarevas et al. employed a sophisticated modeling and research design that was more rigorous than designs used in observational studies. Also, they illustrated the analytic steps they took to rule out alternative interpretations and triangulate their findings, for example examining both state bans and federal bans. They helped build the foundation for future studies while overcoming the limitations of previous research.

MOVING MASS SHOOTING SCIENCE FORWARD

Later research would draw the line in the sand where this study ended or dig into other nuances not addressed.[11] For example, Klarevas et al. included both national and state-level bans on LCMs; however, the national legislation and some states also included a ban on assault weapons, so we cannot say with certainty that it was a ban on LCMs, a ban on assault weapons, or a combination. Because assault weapons often (but not always) include LCMs while other guns that are not assault weapons also include LCMs, it is possible that a ban on either would attenuate mass shootings. In addition, although LCM bans were effective, significant loopholes remained that would-be shooters could get around to access illegal weapons and magazines. Most policies grandfathered in individuals who already owned assault weapons with LCMs. Other research would go on to identify stolen guns as pervasive in homicide shootings, so not removing assault weapons and LCMs from the population might reduce the impact of bans.

Moreover, although the weapon bans were applied to gun sales, private vendors were not subject to the bans. After the federal assault weapon ban sunset in 2004, motivated shooters in states with bans were able to easily travel to states without bans, underscoring the need for national policies. Finally, although Klarevas et al. made a good case for including only mass shootings that resulted in six deaths or more, it is important to know whether LCMs empower mass shooters in general, even in the case of shootings with a lower lethality threshold.

SUMMARY

Klarevas, Conner, and Hemenway published an important study that was not popular in select political circles or among gun manufacturers and the

National Rifle Association. They firmly established that high-lethality mass shootings can be prevented through policies. Their investigation built on previous mass shooting research, and the gun scholars who came afterward used the study to agree or disagree but always to push the knowledge base forward. Scholars cite this seminal study because of its robustness and quality. Louis Klarevas, Andrew Conner, and David Hemenway are agitators who got into what the late, great Representative John Lewis (D, GA) called "good trouble, necessary trouble."[12] Scientists walk away from this study knowing that policies can prevent gun deaths, whereas nonacademic citizens have learned that commonsense policies informed by scientific rigor, such as bans on LCMs, help to prevent public massacres. Finally, researchers have learned that they must persevere, sometimes in hostile environments, to inform injury prevention. AJPH

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L. A. Post was responsible for drafting the outline and for the overview of mass shootings, the Supreme Court ruling, definitions of mass shootings, laws on large-capacity magazines, and assault weapon bans. M. Mason was responsible for the policy-making process, various media citations, and metrics of impact factors.

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Gun Control for Health: A Public Health of Consequence, December 2022

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See also Post and Mason, p. 1707 and Rowhani-Rahbar et al., p. 1783.

On June 23, 2022, the US Supreme Court ruled in *New York State Rifle & Pistol Association (NYSPRA) v. Bruen* that the New York State law requiring individuals to show proper cause to obtain a license to carry a concealed firearm in public places for purposes of self-defense was unconstitutional (https://bit.ly/3DBV4vF). The significance of this ruling from political and public health perspectives cannot be underestimated. Against the backdrop of growing political partisanship among US legislators and in the US Supreme Court, the persistent lobbying by gun rights groups and the gun industry to loosen gun regulations and promote gun sales, as evidenced by the *NYSPRA* ruling, exemplifies how commercial determinants undermine health and well-being. The commercial interests of the gun lobby and the gun industry that limit research and drive laws and practices to sustain the availability and presence of guns in the United States cause immediate and horrific public health harms—mass shootings, mass murders, homicides, suicides, and unintentional gun-related injuries and deaths. The physical and emotional costs of gun-related injuries and deaths to survivors, their friends, and families are staggering. An evaluation funded by Everytown for Gun Safety concluded that gun violence costs Americans $557 billion annually—the bulk of which is attributed to quality-of-life costs for victims and their families ($489.1 billion) and medical costs ($2.8 billion) (https://bit.ly/3Ubhw4N).

PUBLIC HEALTH, NOT CORPORATE HEALTH

In their 2018 commentary, McKee and Stuckler[1] presented a summary of key manifestations of corporate power that influence health. Two of these manifestations—setting the narrative and setting the rules—are clearly part of the playbook of the gun lobby as they seek to dismantle gun control legislation. By focusing on a narrative of gun "rights" in legislative and judicial decisions and pouring money to back politicians who will not support gun control prevention or research efforts, the gun industry has ensured that corporate power supersedes public health (https://bit.ly/2CnxRdo).

The Supreme Court ruling in the *NYSPRA* case is the latest key decision limiting gun control since the decision in *District of Columbia v. Heller* (2008; https://bit.ly/3sqCnnP) and in *McDonald v. City of Chicago* (2010, https://bit.ly/3Fcajfv). The culmination of these decisions demonstrates the steady chipping away at federal gun regulations in the name of upholding the Second Amendment, but, in reality, to sustain the corporate and financial interests of the gun industry in the United States. Quite simply, the gun lobby in the United States is unlike that of any other special interest group. Although others, namely tobacco,[2] alcohol,[3] food,[4] and the sugar-sweetened beverages[5] industries, have been subjected to widespread and growing regulation for marketing products that cause health-related harm, the gun lobby has remained largely unregulated, despite the pervasiveness of gun violence in the United States.

Although many are aware of the Dickey Amendment, few are likely to know that a major impetus for this amendment was a 1993 study by Kellerman et al.[6] showing that the presence of a gun in a home increased the odds of homicide. In an effort to stall robust research on gun violence, the National Rifle Association (NRA) lobbied for the Dickey Amendment to the 1996 US spending bill, an amendment that effectively banned federal funding to the Centers for Disease Control and Prevention (CDC) for research that could be used to advocate or promote gun control.[7] In 2009, Branas et al.[8] reported findings from a National Institute on Alcohol Abuse and Alcoholism–funded study showing that individuals in possession of a gun were four times more likely to be shot in an assault than those not in possession of a gun. In 2012, once again with backing by the NRA, the US omnibus spending bill expanded its ban on federally funded gun control research to

AJPH December 2022, Vol 112, No. 12

include the National Institutes of Health (NIH) as well as the CDC. The presence of this ban for more than 20 years is one of the most prominent examples of lobbying and corporate manifestation of power and of setting rules that eliminate funding for gun control research. Although a small number of researchers were able to continue carrying out gun-related research, the Dickey Amendment essentially eliminated the possibility of creating a robust evidence base on gun violence prevention.

GROWING THE GUN CONTROL EVIDENCE BASE

Recently, the language of the Dickey Amendment has been clarified to allow the CDC and NIH to conduct gun violence–related research, and a $25 million allocation, distributed evenly between the CDC and NIH, was earmarked for gun violence prevention research. These funds provide what amounts to seed funding to conduct research on the impact of federal and state gun legislation compared with funding for other health issues. Despite this slow and small start, more funding and research are critically necessary to establish an evidence base that, it is hoped, can inform the myriad of laws, policies, and practices that will be required to comprehensively limit the availability of and access to guns.

In this issue of *AJPH*, Post and Mason (p. 1707) reflect on the significance of the study by Klarevas et al.[9] conducted during the era of the Dickey Amendment. As Post writes, the contribution of the Klarevas et al. study is significant for providing additional empirical evidence on the effect of state and federal large-capacity magazine (LCM) bans on the frequency and lethality of mass shootings. The study included 69 mass shooting events between 1990 and 2017, when state (enacted in New Jersey in 1990 and still in place in nine states and the District of Columbia) and federal (enacted in 1994, expired in 2004) legislation was in place.[9] Klarevas et al. found that mass shootings where LCMs were used were more likely to have higher fatalities than those where an LCM was not used and that, in states lacking LCM bans, the incidence of high-fatality mass shootings was more than twice that in states with LCM bans.

In the wake of the *NYSPRA* ruling, which opens the door to loosening restrictions on handgun carrying, the study by Rowhani-Rahbar et al. (p. 1783) provides much-needed baseline evidence on trends in handgun carrying in the United States. Based on a nationally representative sample of gun owners, the study found that the number of handgun owners who carried their guns on a monthly basis increased dramatically from 9 million in 2015 to 16 million in 2019 and that daily handgun carrying doubled during this period. Future studies will be needed to understand the links between handgun carrying and involvement in gun violence, whether guns carried are concealed or open, and in what types of public spaces guns are carried. Additionally, studies building on prior work examining how and where guns are safely stored will provide information to inform interventions to prevent suicides as well as unintentional injuries.[10] All of these important questions require careful investigation, and it is hoped they will be supported by future funding.

CONCLUSION

Caught between our national struggle between democratic freedoms and corporate interests are the individual and societal harms inflicted by gun violence. Between January 1 and October 1, 2022, there were 515 mass shootings (shootings of more than four people) and 21 mass murders (murder of four or more people in a mass shooting; https://www.gunviolencearchive.org). During this same period, 15 547 persons were murdered (intentional and unintentional homicide, defensive gun use) and 18 348 persons committed suicide with a gun. In addition to more research on those who already own a gun or will become new gun owners,[11] as well as how they will carry and use guns, parallel efforts to examine the toll of gun violence exposure on individuals and communities, as well as effective prevention, are also necessary. AJPH

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Communities, Mistrust, and Implementation: Addressing a Large Gap in the National Strategy for COVID-19 and Future Pandemics

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The United States is at a turning point in responding to the ongoing pandemic. Although the severity of disease associated with the rapidly evolving severe acute respiratory syndrome coronavirus 2 (SARS-CoV-2) appears to have decreased, the huge number of infections and reinfections are impacting quality of life, disrupting public health and economic and social functioning, and presenting a potentially large post-acute epidemic burden from "long COVID," even while additional variants continue to emerge with unpredictable risks.

NATIONAL STRATEGIES FOR MANAGING CURRENT AND FUTURE PANDEMICS

This has promoted a re-evaluation of national strategies for managing the current and future pandemics. Commentators such as Emanuel et al.[1] have proposed focusing on appropriate risk thresholds; rebuilding the public health system; improving testing, disease, and genomic surveillance systems; building ventilation systems and personal protective devices; investing in next-generation vaccines; and accelerating the development of antiviral treatments. Coupled with the recent announcement of grand initiatives being undertaken, such as the $150 million Pandemic Prevention Institute funded by the Rockefeller Foundation to support global data collection and the $500 million Center for Forecasting and Outbreak Analytics being created by the US Centers for Disease Control and Prevention (CDC),[2] a consensus is emerging on the broad outlines of a comprehensive response to the current and future pandemics.

ADDRESSING MISTRUST

But does this vision adequately address fundamental aspects of pandemics? Although many have acknowledged the challenge of public distrust of health agencies and evidence-based policies resulting in a lack of adherence to risk-mitigation measures, a national strategy for addressing such distrust is lacking. This is a glaring omission, considering that, for example, vaccine hesitancy and denial are likely responsible for a large portion of the estimated 319 000 excess COVID-19 deaths that vaccinations could have been prevented in the United States (as of April 30, 2022[3]) since vaccines became widely available.

In terms of those affected, national surveys have shown that skepticism and vaccine hesitancy are strongly associated with Republican political preferences and conservative religious beliefs.[4] As another example, distrust of institutions (based in large part on a long history of racism related to health care and medical research) leading to low rates of vaccination has been shown to be likely responsible for the disproportionate impact of COVID-19 on African Americans,[5] one of the many health inequities made starkly apparent by the pandemic.

The challenge posed by distrust, of course, is complex, as it relates to some segments of the population with respect to institutions, political parties, scientific experts, and media outlets. Distrust has also been fueled by a lack of clear communication about the need to change policies as scientific information evolved, misinformation, uncertainty about the content or sources of information, and contradictory information. Surveys tracking public attitudes found that 78% of adults say they have heard at least one of eight different false statements about COVID-19 that they believe to be true or they are unsure of its veracity.[6]

We propose that a national strategy is essential to address distrust as a critical factor in controlling pandemics and will require attention and investments in several aspects of "population health implementation science," an area that remains consistently underrecognized, underfunded, and understudied.[7] These include, for example, risk communication methods; the epidemiology of information and disinformation; the impact on attitudes and behaviors of popular media, social media, and other forms of communication; controlled trials of policy and messaging interventions; and the direct involvement of communities as sources of vital information and participants in the planning and conduct of research. In addition, a national strategy is needed to accommodate regional differences, including the testing and adoption of optimal strategies that may differ widely within and between populations and regions in terms of racism and ethnicity, culture, socioeconomic status, urban versus rural, levels of education, gender identity, and other factors.

COMMUNITY-BASED PARTICIPATORY RESEARCH AND EDUCATION

We describe an example that was implemented first in Los Angeles, CA, during April 2021 to April 2022, and then in 34 cities across the country. The Vaccinate LA campaign is a joint effort by 14 units of the University of Southern California (USC) and Children's Hospital Los Angeles, two local creative agencies: Wondros, and Everyone Can Eat Productions, more than 160 community-based organizations, the USC Keck School of Medicine Stay Connected LA program, and a community advisory board.[8] The campaign implemented a mass media educational effort focused on Black and Latino/a/x populations, developed and deployed trainings of community vaccine navigators, and assessed the impact on attitudes, beliefs, and behaviors toward COVID-19 vaccinations.

The goal was to address pervasive misinformation and distrust at the community level and provide access to COVID-19 vaccines in 34 zip codes in the eastern and south–central areas of Los Angeles experiencing low vaccination rates, high hospitalizations, and deaths. Using a community-based participatory research and education approach,[9] the program incorporated listening sessions (focus groups and town-hall meetings), information delivery, interactive and field-based activities (pop-up vaccination clinics or sites), and social media. Community vaccine navigators (*promotores de salud* and community health workers) who speak Spanish and English were trained and deployed door-to-door, at community events, and at pop-up vaccination clinics, and they provided one-on-one counseling to address misinformation, increase trust, respond to frequently asked questions regarding fears and concerns of vaccinations, and link individuals directly to vaccines (M. Kipke, August 2022, Vaccinate LA Final Report to the W. M. Keck Foundation).

Forty-two focus groups with more than 300 participants and 21 town hall meetings with more than 200 participants informed the campaign. These and other baseline and posttest data, changes in frequently asked questions, social media data analytics, and surveys revealed changes in attitudes and beliefs regarding vaccinations. In early 2022, the program's success led to its adoption and adaptation by the National Alliance for Hispanic Health, an organization serving more than 15 million Hispanics nationwide, which, in turn, resulted in the training of 450 community vaccine navigators. Training included updated COVID-19 information, approaches to handling misinformation and frequently asked questions, debunking myths and addressing conspiracy theories, and the use of evidence-based approaches and innovative multimedia strategies including culturally adapted films ("Granny's Birthday," "Of Reasons and Rumors") developed by local Latino/a/x and Black filmmakers.

A digital communication campaign was conducted with Hollywood Health and Society, producing "Life Noggin," an animated science show on YouTube reaching 3.26 million viewers[10]; postings for social media, and production of 41 #ShareYourWhy videos, resulting in 2.9 million views. A *fotonovela* in Spanish was produced and disseminated through newspapers, including *La Opinion*, with a Spanish-language readership of 540 000. Pop-up events were conducted with community partners to get shots in arms and supported local artists in an art-meets–public health program (Stay Connected Los Angeles).[11]

At a total cost of $1.2 million along with efforts coordinated with the Los Angeles County Department of Public Health, these community-based participatory research and education approaches resulted in a vaccination rate that was 30% higher than predicted in the targeted areas based on Los Angeles County averages. Nationally, with support from the CDC, trained vaccine navigators conducted one million community vaccine navigator consultations across the United States resulting in 500 000 shots in arms in 38 cities nationwide. While promising, limitations that became apparent included unclear generalizability to Blacks across the country, the need to stay current with an ever-changing virus, the wide spread of the virus across multiple geographic areas, and the need to combat a constant stream of misinformation.

COMMUNITY-BASED IMPLEMENTATION SCIENCE

This experience and a handful of other examples[12,13] align with emerging principles of population health implementation science that must become part of national strategy to address pandemics. Such a strategy should include national funding of research as well as regional centers of excellence that can develop and evaluate approaches tailored to specific communities, with engaged community participation in the design and implementation being an essential component. Such strategies need not be expensive; in fact, public health implementation strategies developed in resource-poor countries have been shown to be of value in rich countries, an example of reverse innovation.[14] Most importantly, programs that are developed with and for communities, utilizing community-based and participatory principles, can reach populations with greater accuracy and effectiveness and provide a trusted source of information for sound community, familial, and individual decision-making. AJPH

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Knowledge and Practice of Pinworm Infection in Preschool Children, Jiangsu Province, China, 2019–2020

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We conducted a two-year (2019–2020) longitudinal study in Jiangsu Province, China to analyze risk factors of pinworm infection and evaluate the effect of behavior change communication–based (BCC-based) interventions in preschool children. The positive rate of pinworm infection was higher in private preschool (2%) than in public preschool (0.24%). Poor sanitation behaviors were risk factors among private preschool children. BCC-based intervention could improve knowledge and practice and reduce pinworm infection. This study may help fill in gaps in pinworm control. (*Am J Public Health*. 2022;112(12):1716–1720. https://doi.org/10.2105/AJPH.2022.307067)

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Enterobiasis, caused by pinworm (*Enterobius vermicularis*) infection, is one of the most prevalent parasitic diseases among children regardless of their socioeconomic level, culture, or race.[1–3] In severe cases, insomnia, weight loss, vomiting, abdominal pain, and appendicitis can appear. Pinworm eggs are transmitted from person to person, directly via anus-to-mouth contamination, finger contamination, or through indirect touch of contaminated objects (e.g., toys and classroom tables). Preventing infection and reinfection may be challenging because of a simple life cycle.[4] Children with poor personal hygiene are susceptible to pinworm infection and reinfection, especially those in crowded organizations.[5] However, enterobiasis in children is considerably neglected by parents and health officials. Enterobiasis is rarely a subject of in-depth epidemiological inquiries, in developed or developing countries, despite its wide occurrence. Screening of key populations, analyzing risk factors, and precise interventions are conducive and necessary to pinworm control. Behavior change communication (BCC) is widely used to promote and sustain healthy changes in behavior through tailored health messages and approaches.[6] BCC-based intervention may facilitate pinworm control at the community or school level.

INTERVENTION AND IMPLEMENTATION

We conducted two preschool-level surveys to implement data collection. Letters of information, informed consent forms, and questionnaires were given to parents or principal caretakers prior to the survey. The BCC-based intervention included health education and providing health consultation services. Health education comprised guiding daily hygiene, developing hygiene habits, holding lectures, distributing leaflets, and providing health consultation services. We recruited experts from municipal-level and province-level Centers for Disease Control and Prevention to provide health consultation services. The parents received feedback on the results of pinworm detection. We told parents of children testing positive to seek medical treatment and advice.

We treated Gulou District and Gangzha District as intervention groups and Haimen District and Guangling District as control groups. The inclusive and exclusive criteria are shown in panel b of Figure B (available as a supplement to the online version of this article at https://www.ajph.org). However, we excluded the Haimen District children from the control group because of the preschool adjustment made by the local government at the beginning of 2020.

BCC-based health education was implemented with the aim of having teachers and parents cultivate conscious hygiene habits in the children studied. Lectures were held four times per year (twice a semester) by health workers. Leaflets were given out to children and parents every semester (two semesters a year). Moreover, teachers in these preschools were requested to hold a lecture about pinworm infections and prevention at the parents' meeting.

Considering the low pinworm infection rates in public preschool, we conducted BCC-based intervention in private preschools. Sample collection was performed in the morning, before the children defecated and bathed. Health workers from county- and municipal-level Centers for Disease Control and Prevention took one sample from each child. The adhesive cellophane tape swab (patent number: ZL201420707045.8) was used over the perianal skin and was then inspected by a trained municipal-level microbiologist and checked by an expert from Jiangsu Institute of Parasitic Diseases. One or more eggs found under the microscope indicated a pinworm infection. We invited the child's parents or principal caretakers to complete a questionnaire (Table A, available as a supplement to the online version of this article at https://www.ajph.org); questions included information related to demographics, household sanitary conditions, and knowledge and practice regarding enterobiasis. The content of the questionnaire was the same at baseline and at follow-up.

We used SPSS version 13.0 (SPSS Inc, Chicago, IL) to conduct data analyses. We used the χ^2 test to test significant differences in group outcomes. We conducted multivariate logistic regression by treating the grouping variable (intervention group or control group) as outcome and gender, age, residence, mother's educational level, type of flooring in the home, and hygiene habits among the follow-up population as covariates to achieve the propensity score, which was classified into four groups according to quartiles. We applied the Cochran–Mantel–Haenszel test, adjusted by grouped propensity score, to test statistical differences between the intervention group and control group. We estimated odds ratios and 95% confidence intervals using the Poisson loglinear model of risk factors for pinworm infection at baseline (2019). We applied principal component analysis to detect hygiene factors. The factor of the principal component analysis with an eigenvalue above 1 was retained. A *P* value of less than .05 indicated a statistically significant difference.

PLACE, TIME, AND PERSONS

The study included preschool children in six districts of Nanjing, Yangzhou, Nantong, and Yancheng Prefectures (Figure A, available as a supplement to the online version of this article at http://www.ajph.org). The program (Figure B) was conducted September 1, 2019 through October 31, 2020 in Jiangsu Province, China. We made epidemiological assessments and analyzed associated risk factors of pinworm infection after a cross-sectional survey (September–October 2019). We then implemented a BCC-based intervention. The participants were followed up from September to October 2020.

Preschool children aged two to six years, as well as their parents or principal caretakers, were included. In each study site, we selected two types of preschools: public preschools, which only admit children from permanent population families, and private preschools, which only admit children from transient population families. A majority of children came from one-child families. If a family had multiple children aged two to six years attending the selected preschool, all children were recruited as respondents.

PURPOSE

Although enterobiasis was mostly controlled, it was not completely eliminated.[7,8] The aim of this study was to explore risk factors and BCC-based intervention approach for pinworm infection in Jiangsu Province, China.

EVALUATION AND ADVERSE EFFECTS

A total of 3678 preschool children (1697 from public preschool and 1981 from private preschool) aged two to six years were enrolled in 2019 (Table 1). The overall rate of pinworm infection was 1.2%. At baseline (in 2019), 54% were boys, and the mean (±SD) age of the children was 4.4 (±1) years; a majority of children were aged four to five years (65.6% and 64.0% for private preschool and public preschool, respectively); four (0.24%) and 41 (2%) positive cases of pinworm infection were found in public preschools and private preschools, respectively. Hygiene behaviors of *Enterobius vermicularis* infection among preschool children are shown in Table B (available as a supplement to the online version of this article at https://www.ajph.org).

Improvement of Knowledge and Practice

Knowledge improved in the intervention group, whereas it decreased in the control group (Table 2). Moreover, private preschool children showed greater improvement in behavior after BCC-based intervention, especially in relationship to washing hands, sucking fingers, and sucking toys and pens (Table C,

TABLE 1— Baseline Demographic Characteristics of Preschool Children in *Enterobius vermicularis* Infection Intervention: Jiangsu Province, China, 2019

	Private Preschool			Public Preschool		
	No. (%)	No. of Positive Cases (Infection Rate, %)	P[a]	No. (%)	No. of Positive Cases (Infection Rate, %)	P[a]
Gender			.75			.63
Male	1077 (54.4)	21 (1.95)		923 (54.4)	3 (0.33)	
Female	904 (45.6)	20 (2.21)		774 (45.6)	1 (0.13)	
Age, y			< .001			.99
2	8 (0.4)	0 (0.00)		7 (0.4)	0 (0.00)	
3	398 (20.1)	8 (2.01)		365 (21.5)	1 (0.27)	
4	614 (31.0)	2 (0.33)		528 (31.1)	1 (0.19)	
5	686 (34.6)	18 (2.62)		557 (32.9)	2 (0.36)	
6	275 (13.9)	13 (4.73)		240 (14.1)	0 (0.00)	
Grade			.04			.99
Bottom	554 (28.0)	8 (1.44)		533 (31.4)	1 (0.19)	
Middle	783 (39.5)	12 (1.53)		542 (31.9)	1 (0.18)	
Top	644 (32.5)	21 (3.26)		622 (36.7)	2 (0.32)	
Residence			.99			.58
Urban	1001 (50.5)	21 (2.10)		1224 (72.1)	4 (0.33)	
Rural	980 (49.5)	20 (2.04)		473 (27.9)	0 (0.00)	
Mother's educational level			.52			.052
Primary school or lower	100 (5.0)	2 (2.00)		34 (1.9)	1 (2.94)	
Secondary school	1657 (83.6)	37 (2.23)		1171 (69.0)	3 (0.26)	
Diploma, bachelor's, or higher	224 (11.3)	2 (0.89)		492 (29.0)	0 (0.00)	
Family income level			.023			.001
Low	549 (27.7)	18 (3.28)		236 (13.9)	4 (1.69)	
Medium	1414 (71.4)	22 (1.56)		1447 (85.3)	0 (0.00)	
High	18 (0.9)	1 (5.56)		14 (0.8)	0 (0.00)	
Type of house floor			.056			.21
Brick or wood	1494 (75.4)	26 (1.74)		1598 (94.2)	3 (0.19)	
Cement	457 (23.1)	13 (2.84)		92 (5.4)	1 (1.09)	
Soil	30 (1.5)	2 (6.67)		7 (0.4)	0 (0.00)	
Total	1981	41 (2.01)	NA	1697	4 (0.24)	NA

Note. NA = not applicable.
[a]By the Fisher exact test.

available as a supplement to the online version of this article at https://www.ajph.org).

Reduced Positive Rate, But New Infections

The intervention group consisted of 723 children from Gulou and Gangzha Districts (positive rate: 4.3%), whereas the control group comprised 258 children from Gangling District (positive rate: 0.4%; panel b, Figure B).

In 2020, we followed up 740 children from baseline: 505 children in the intervention group (51.9% boys; 60.2% of children aged four to five years) and 235 children in the control group (50.6% boys; 67.7% of children aged four to five years). Following the one-year intervention, 18 of the children positive at baseline were found to be negative; however, seven new infections were found (positive rate: 1.4%). There was no infection in the control group during follow-up.

TABLE 2— Knowledge Awareness Rate of Enterobiasis at Baseline and Follow-Up Among Parents of Private Preschool Children: Jiangsu Province, China, 2019–2020

	Intervention Group			Control Group			Intervention – Control	
Parental Knowledge	Baseline (n = 723), %	Follow-Up (n = 505), %	Increment, Percentage Points	Baseline (n = 258)	Follow-Up (n = 235)	Increment, Percentage Points	Percentage Point Difference	P^a
Knows enterobiasis	58	64	+6	61	44	−17	+20	< .001
Knows it is contagious	62	62	+0	58	58	−0	+4	.3
Knows its route of infection	53	53	+0	61	52	−9	+1	.91
Knows its main parasitic site	47	44	−3	56	48	−8	−4	.04
Knows its symptoms	49	55	+6	53	43	−10	+12	.002
Knows its susceptible population	84	85	+1	85	86	+1	−1	.58
Knows drug used for treatment	45	49	+4	37	33	−4	+16	< .001
Knows prophylactic measures	81	84	+3	80	77	−3	+7	.04

[a]By the Cochran−Mantel−Haenszel test.

Factors Associated With Infections

We used principal component analysis to develop hygiene factors from nine variables related to personal hygiene behaviors (Table D, available as a supplement to the online version of this article at https://www.ajph.org). We retained four principal components (PCs). PC1 indicated a composite factor of sucking habit; PC2 indicated the habit factor of washing hands; PC3 indicated a composite factor of maintaining personal hygiene and tidiness; PC4 indicated bathing habits. According to the Poisson loglinear model of children from private preschool, risk factors found to be associated with pinworm infections were age (odds ratio [OR] = 1.6; 95% confidence interval [CI] = 1.2, 2.3), PC3 (OR = 1.2; 95% CI = 1.01, 1.5), and PC4 (OR = 1.4, 95% CI = 1.1, 1.8; Table E, available as a supplement to the online version of this article at https://www.ajph.org). Among public preschool children, only brick or wood floor (reference = cement floor) resulted as a risk factor (OR = 19.9; 95% CI = 13.1, 30.4; Table E).

We observed no adverse effects.

SUSTAINABILITY

Children's pinworm infections have been overlooked because there is a serious lack of studies on the topic. This work calls for more attention to be paid to pinworm infections among preschool children, as well as for the sustainable practice of pinworm control. Moreover, in this study, we have developed a protocol of BCC-based intervention approach and control strategy, incorporated in annual parasitological surveys with funding support, which may facilitate the formation and continuity of best practices for pinworm infection control in preschool.

PUBLIC HEALTH SIGNIFICANCE

To the best of our knowledge, the present study is the first report on the prevalence of pinworm infections among private and public preschool children. To better understand the context of this study, it's important to note that the distinction between public and private schools may have different connotations in China compared with other parts of the world. This study provides an in-depth and new insight into preschool-based pinworm risk and intervention efforts worldwide. BCC-based intervention, which could improve knowledge and practice and reduce pinworm infection, could be further applied for pinworm control among children, especially private preschool children. AJPH

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CONTRIBUTORS

J. Cao and Y. Dai conceptualized the study. F. Mao, Y. Yang, and Q. Zhang analyzed the data, drafted the manuscript, and contributed to data analysis and interpretation. X. Ding, X. Xu, and Y. Chen assisted with data collection. All authors critically reviewed and revised the manuscript.

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CONFLICTS OF INTEREST

The authors have no conflicts of interest to declare.

HUMAN PARTICIPANT PROTECTION

The present study was approved by the institutional review board of Jiangsu Institute of Parasitic Diseases (JIPD-2018-002), Wuxi, China.

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Community-Based COVID-19 Vaccine Clinics in Medically Underserved Neighborhoods to Improve Access and Equity, Philadelphia, 2021–2022

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Vaccination remains key to reducing the risk of COVID-19–related severe illness and death. Because of historic medical exclusion and barriers to access, Black communities have had lower rates of COVID-19 vaccination than White communities. We describe the efforts of an academic medical institution to implement community-based COVID-19 vaccine clinics in medically underserved neighborhoods in Philadelphia, Pennsylvania. Over a 13-month period (April 2021–April 2022), the initiative delivered 9038 vaccine doses to community members, a majority of whom (57%) identified as Black. (*Am J Public Health*. 2022;112(12):1721–1725. https://doi.org/10.2105/AJPH.2022.307030)

AJPH December 2022, Vol 112, No. 12

To improve COVID-19 vaccine access among medically underserved and vulnerable populations in Philadelphia, Pennsylvania, we implemented low-barrier vaccine clinics throughout Philadelphia, in collaboration with the Philadelphia Department of Public Health, the School District of Philadelphia, Philadelphia Parks and Recreation, faith-based institutions, community organizations, and professional sports organizations.

INTERVENTION AND IMPLEMENTATION

The University of Pennsylvania and the University of Pennsylvania Health Systems hosted large-scale COVID-19 vaccination clinics for Philadelphia residents in February 2021.[1] In April 2021, when vaccine eligibility was expanded to include anyone aged 16 years or older, the Department of Family Medicine and Community Health began implementing community-based pop-up clinics in West and Southwest Philadelphia.

PLACE, TIME, AND PERSONS

The clinics targeted communities of color that faced financial and geographic barriers to vaccine access through health care centers and retail pharmacies and were primarily located in neighborhoods with low COVID-19 vaccination rates. From April 2021 to April 2022, we hosted 68 clinics in trusted neighborhood venues at the request of organizations with deep community ties. We typically aimed to host two clinics at three-week intervals to provide first and second doses. Our community-based hospital also maintained walk-in vaccine access five days a week.

PURPOSE

More than one third (33.7%) of the population of West Philadelphia is living in poverty, compared with 10.5% nationally.[2] For many reasons (e.g., historic exclusion as a result of systemic racism, geographic barriers to access), Black adults and children have had lower COVID-19 vaccination rates than have those in White communities.[3–5] The goal of this program was to implement frequent, low-barrier COVID-19 vaccine clinics in West and Southwest Philadelphia. We also aimed to promote patient choice by offering all available COVID-19 vaccinations (vs earlier mass vaccination efforts that typically offered single manufacturer vaccines).

Planning and Registration

We identified clinic locations through community partner requests and included K–12 schools, recreation centers, restaurants, religious institutions, and youth and athletic organizations. Requests exceeded our capacity to host clinics, so we prioritized locations that were accessible by public transit, were in low-vaccination neighborhoods, and had large indoor spaces to facilitate physical distancing. People could preregister via a text message–based system or walk in without appointments.[1]

Recruitment strategies included School District of Philadelphia–initiated robocalls and digital communications; SMS (short message service) campaigns in which all individuals who had previously registered for a vaccine clinic received a message in advance of our next clinic encouraging them to refer individuals for vaccination; physical and digital flyers shared with community partners; and virtual town halls with clinicians to answer questions.

Staffing

Clinics were staffed by volunteers who were recruited through listservs and personal outreach; volunteers signed up using a Web-based platform. Nonclinical volunteer roles included two operations leaders and five to 10 members of support staff (e.g., clinic navigation and registration). Clinical volunteers were Pennsylvania-licensed physicians, nurses, advanced practice providers, or pharmacists filling the following roles: one medical director, three to five vaccine preparation specialists, two to four postvaccine monitors, and five to 15 vaccinators. All vaccinators were required to complete a 10-minute Web-based training session before their first clinic. With clinical supervision, medical, dental, and pharmacy students also served as vaccinators. More than 500 staff members volunteered at 68 clinics, approximately 60% of whom were nonclinical and 40% clinicians.

Logistics

Clinics offered all COVID-19 vaccines approved formally or under US Food and Drug Administration emergency use authorization. Given the multiple manufacturers and doses available, we designed a color-coded system with safety protocols that included just-in-time training and multiple built-in color-coded checkpoints to ensure that patients received the correct vaccine (Figure 1). Upon entry to the vaccination clinic, patients were assigned color-coded paperwork indicating their designated vaccine:

1. Pfizer Blue—Pfizer-BioNTech 0.3-milliliter (mL) dose for those aged 12 years and older (primary series and booster doses);
2. Pfizer Orange—Pfizer-BioNTech 0.2-mL dose for those aged five to 11 years;
3. Moderna Green—Moderna 0.5-mL dose for primary series;
4. Moderna Pink—Moderna 0.25-mL dose for booster (as of October 2021 approval); or
5. JNJ Yellow—Johnson & Johnson/Janssen (primary series and booster doses).

Visual, written, and verbal communication all used the color-coded names. Vaccine storage, syringes, and labels were all similarly coded.

At check-in, patients received intake documents, including an emergency use authorization information packet, paper registration, and consent documents for patients younger than 18 years. Parents or guardians were required to accompany children aged five to 14 years receiving primary series and children aged 12 to 17 years receiving a booster. Parents or guardians of children aged 15 to 17 years receiving primary series were able to provide consent via telephone with an on-site physician.

Although the Philadelphia Department of Public Health allows minors aged 12 years and older to consent to their own COVID-19 immunization without the consent of a parent or guardian under an emergency use authorization, we included parent or guardian consent in our processes to prioritize community trust.[6] After check-in, staff completed patient registration using a Web-based system, also serving as a second safety checkpoint to ensure that patients were assigned the correct vaccine. Because of the collection of protected health information, all registration and data input was completed on a secure, portable network connecting to a remote server.

Once check-in and registration were complete, patients were directed to a vaccine station. Clinicians displayed a "READY" sign to let volunteers know they were available to vaccinate the next patient. Vaccine vials were held in baskets of the assigned corresponding color. Before vaccination, clinicians completed a final confirmation of vaccine type with color-coded syringes. After vaccine administration, patients were directed to complete 15 to 30 minutes of clinical observation. Patients were not permitted to leave the clinic until the observer collected their registration paper and documented the time of observation completion.

EVALUATION AND ADVERSE EFFECTS

From April 2021 to April 2022, our team vaccinated 9038 patients across

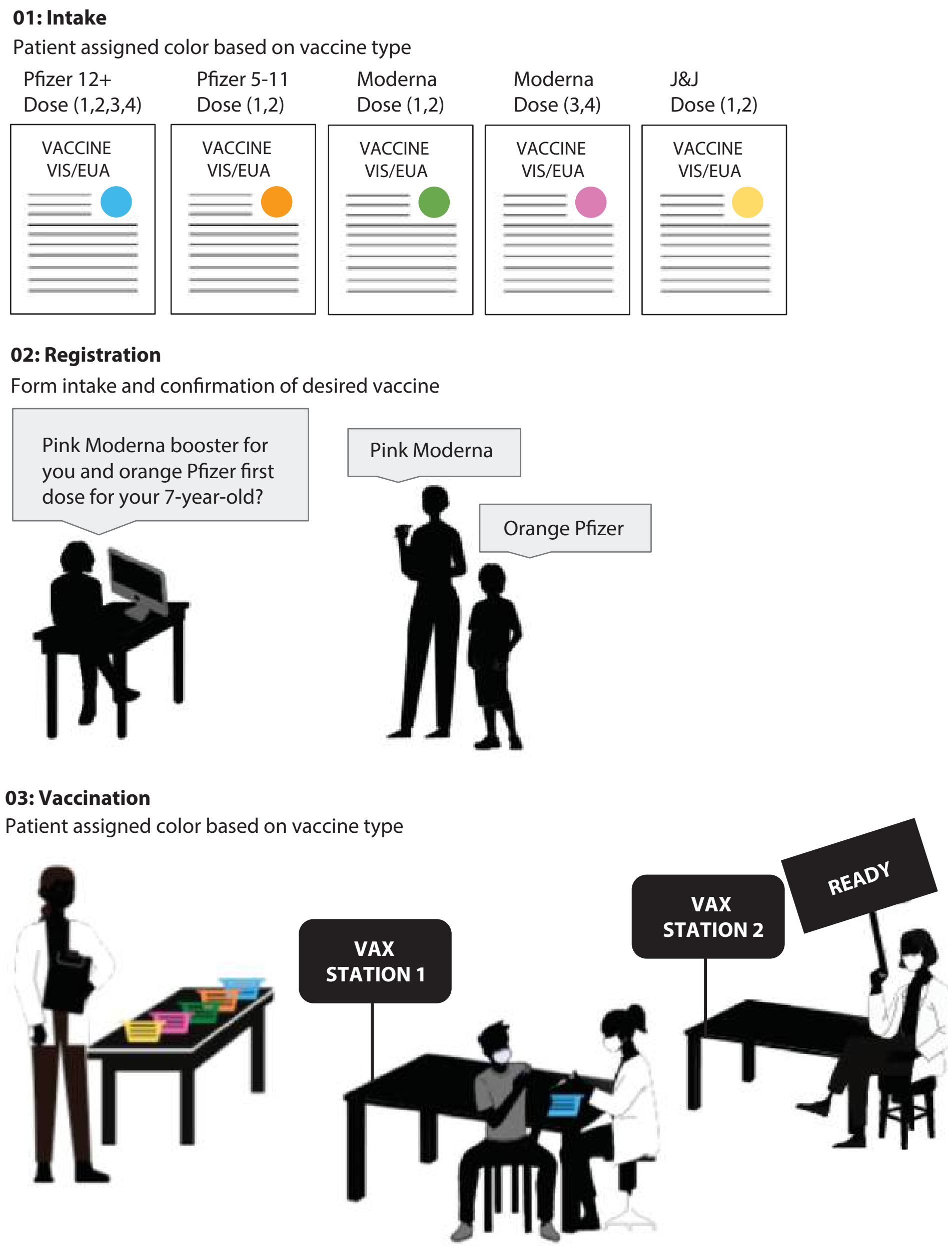

FIGURE 1— Clinic Color-Coded Safety System for Penn Medicine's Pop-up COVID-19 Vaccine Clinics: Philadelphia, PA, April 2021–April 2022

Note. EUA = emergency use authorization; J&J = Johnson & Johnson; VAX = vaccination; VIS = vaccine information statement.
[a]Colored stickers were placed on a clinic form for collecting basic demographic and health information (e.g., known allergies).

68 clinics (Table 1). Most patients were Black/African American (57%), followed by White (23%), and Asian (7%); 59% of patients were aged 19 to 64 years, and nearly one quarter (24%) were aged five to 18 years. In the same period, the proportion of fully vaccinated residents in our eight target zip codes increased from 16% (50 627) to 62% (196 343) of the population—a 288% increase.[7] We cannot attribute this total gain to our clinics because there were other vaccine providers in the area (e.g., select retail pharmacies). Nonetheless, the 9038 doses delivered—and unenumerated vaccine counseling—provided at our clinics for underserved populations contributed to the overall increase.

A central challenge was ensuring that patients received the correct vaccine, which we addressed using a systems design model to develop the color-coded checkpoint system described previously. Logistical challenges

TABLE 1— Self-Reported Characteristics of Penn Medicine Community COVID-19 Vaccine Clinic Participants: Philadelphia, PA, April 2021– April 2022

Characteristic	No. (%)
Age, y	
5–11	1329 (15)
12–18	811 (9)
19–64	5375 (59)
≥ 65	1403 (16)
Not reported	120 (1)
Race	
Black	5124 (57)
White	2054 (23)
Asian	652 (7)
Other	692 (8)
Not reported	516 (6)
Gender	
Female	4644 (51)
Male	4243 (47)
Other	63 (1)
Not reported	88 (1)

Note. The overall sample (n = 9038) includes data from the 68 community-based vaccine clinics (n = 6343) as well as walk-in clinics hosted at our community hospital (n = 2695).

included securing safe facility spaces that could accessibly accommodate high participant volumes with physical distancing. Other challenges included the physical setup and breakdown of a mobile clinic model that could scale to accommodate up to 500 vaccinations. Finally, the unknown sustainability of and ultimately end of funding from the federal government in March 2022 limited our reach.

SUSTAINABILITY

Our experience facilitating mobile, pop-up, community-based clinics could be adapted for other types of public health interventions, such as flu vaccination and school attendance–mandated immunizations. Color coding from registration limited administration errors and facilitated flow. These efforts, however, are only sustainable with appropriate funding, trustworthy community engagement, and institutional support.

PUBLIC HEALTH SIGNIFICANCE

Vaccination remains a key strategy to stem the tide of the COVID-19 pandemic. More than one year after emergency use authorization approval for vaccination among those aged 12 to 15 years and more than six months after emergency use authorization approval for those aged five to 11 years, vaccine uptake among children remains low. The implementation of centrally located community clinics at trusted venues such as public schools and recreation centers may reduce barriers to COVID-19 vaccination among medically underserved populations as well as children. AJPH

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The Cruel Public Health Consequences of Anti-Immigrant Rhetoric

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See also Miller et al., p. 1738 and Wang et al., p. 1747.

The United States has long used immigration policy to shape the demographic and economic future of the nation.[1] Federal policy balances several priorities, including economic stability, humanitarian goals, family (re)unification, and national security. Changes in immigration policy reflect shifts in the relative emphasis placed on each of these priorities. Since the 1980s, Americans' increasingly polarized views on immigration have contributed to Congress's failure to pass comprehensive immigration reform, frequent changes in aspects of immigration policy that can be regulated without Congress, and a system that is increasingly difficult for immigrants to navigate.[1]

Two articles in this issue of *AJPH* address one federal immigration policy: the public charge rule. The public charge rule is designed to ensure that immigrants who enter the United States will be able to sustain themselves without relying on the government for financial support.[2] In 1999, the public charge rule stated that noncitizens may be denied a green card if they have received general cash assistance or long-term institutionalization funded by the US government or a state, regional, local, or tribal government.[2] Immigrants' use of noncash benefits such as Medicaid and certain cash benefits such as childcare subsidies did not impact their green card eligibility.[2]

In 2017, the Trump administration leaked a draft of a new rule, stating that Medicaid, the Supplemental Nutrition Assistance Program (SNAP), and housing, energy, and childcare assistance would now factor into public charge determinations. The final version of the rule was published in August 2019. The rule was challenged in court and was in effect off and on from October 2019 through March 2021.[3] On September 9, 2022, the US Department of Homeland Security (DHS) released a new version of the public charge rule under which public charge determinations are again based on the guidelines used before 2019.[2]

The Migration Policy Institute estimates that, even under the broad 2019 rule, less than 1% of the 22.1 million noncitizens currently living in the United States could be denied a green card because of public benefits enrollment. Few noncitizens are both subject to the public charge rule and eligible for public benefits. Use of benefits by US citizen children or other household members does not count against a green card applicant in public charge determinations.[2]

Although very few immigrants are subject to the intended effects of this rule, there are widespread unintended effects.[2] As two articles in this issue of *AJPH* show, the 2019 public charge rule led many immigrants to avoid public benefits, even before the rule was implemented. Miller et al. (p. 1738) use the Survey of Income and Program Participation to show that mixed citizenship status households were less likely to use SNAP and school breakfast and lunch programs after the draft rule was leaked in January 2017. Using New York State Medicaid claims, Wang et al. (p. 1747) show that, compared with noncitizens who gave birth in 2014–2016, noncitizens who gave birth after January 2017 enrolled in Medicaid later in pregnancy; were more likely to have delayed, inadequate, or no prenatal care; and had smaller babies.

These findings contribute to growing evidence that after the 2019 rule was announced, enrollment in many means-tested benefits programs declined; immigrants avoided nongovernmental services, including those designed for survivors of domestic and sexual violence; and immigrants were afraid to access COVID-19 testing and vaccination.[2] Older adults, immigrants with disabilities, and US citizen children with immigrant parents were disproportionately impacted.[2] Through the 2022 rule, DHS attempts to limit chilling effects while adhering to the congressional mandate to identify immigrants who are likely to become a public charge.[2] However, experts expect some level of chilling effect to continue.[3]

CHANGING POLICIES CREATE CONFUSION AND MISTRUST

The public charge rule illustrates a broader issue: In the absence of congressional immigration reform,

AJPH December 2022, Vol 112, No. 12

the executive branch is the primary driver of immigration policy. The federal immigration policy landscape changes drastically with each presidential administration. For example, President Obama used executive actions to establish the Deferred Action for Childhood Arrivals (DACA) program and to focus immigration enforcement primarily on immigrants who pose a threat to public safety and national security.[4,5] President Trump issued over 400 executive actions on immigration, including large cuts in refugee resettlement, an attempt to end DACA, and expanded immigration enforcement at the border and within the United States.[4,6] The Biden administration has used executive actions to undo some of President Trump's policies, with varying success.[4,5]

The legislative branch also has a key role in determining immigration policy, because many executive actions have been challenged in court. In 2020–2021, court decisions repeatedly enjoined the 2019 public charge rule, then allowed it to go back into effect.[2] The same is true of other policies, including the termination of DACA and requirements that asylum seekers remain in Mexico while waiting for asylum hearings.[4,6]

Frequent policy changes create confusion, misinformation, and mistrust among immigrants. It is difficult for immigrants and immigrant-serving providers to keep up with policy changes, and it is not always clear how policies will be implemented.[3] Misinformation about the public charge rule includes the belief that noncitizen parents may be denied a green card or even deported if their citizen children enroll in Medicaid, as well as fears that if a green card holder accesses public benefits, they may be ineligible for naturalization.[2,3]

Dramatic shifts in immigration policy also communicate to noncitizens that their presence in the United States is dependent on the whims of the current president. Immigrants act on the basis of not only current policy but also concerns that policies may become more punitive in the future.[3] Even when immigrants know that they or their children are eligible for public benefits and can enroll without endangering their legal status, many decide that it is not worth the risk that these policies may change again soon.[3]

IMPLICATIONS FOR PUBLIC HEALTH RESEARCH

The articles published in this issue of *AJPH* advance our understanding of immigration policy in two key ways. Both Miller et al. and Wang et al. show that chilling effects on food assistance, Medicaid, and prenatal care emerged as soon as the 2019 public charge rule leaked—over two years before it went into effect. Past research has focused on the effects of policies that are passed and implemented[7]; these findings suggest that immigrant health is also harmed by policies that are proposed but fail to pass. Both studies also find that effects of the public charge rule varied on the basis of where immigrants lived. By examining how local context limits or amplifies the effects of federal policies, researchers may identify ways local communities can advance health equity for immigrants.

IMPLICATIONS FOR PUBLIC HEALTH PRACTICE

By deterring immigrants from seeking public benefits and health care, the 2019 public charge rule may have exacerbated the COVID-19 crisis.[2] The 2022 final rule is an important step toward addressing the public health consequences of the 2019 rule, but it must be accompanied by outreach so that immigrants feel safer accessing public benefits.[3] DHS should disseminate information through community-based organizations that have already established trust in immigrant communities,[8] and medical-legal partnerships could incorporate immigration lawyers who can provide up-to-date guidance on changing policies.[3] However, until Congress passes comprehensive immigration reform, public health professionals will face an uphill battle against the misinformation, confusion, mistrust, and fear that currently constrain immigrants' access to health care and public benefits.[3] AJPH

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State-Level Legislation During the COVID-19 Pandemic to Offset the Exclusion of Undocumented Immigrants From Federal Relief Efforts

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See also DeGarmo et al., p. 1757.

The COVID-19 pandemic has worsened the health inequities faced by immigrants, particularly those who are undocumented. Early studies have shown that COVID-19 has disproportionately affected immigrants and their communities.[1] One of the factors that explains the disproportionate morbidity and mortality among immigrants is labor market participation. Undocumented immigrants are predominantly in low-income groups and are uninsured workers who continued to work on-site during the COVID-19 pandemic. Approximately three-fourths of undocumented immigrants work in industries classified as "essential," including agriculture, meatpacking, and construction, among others.[2] Despite the critical participation of undocumented immigrants in essential economic activities, and with the related elevated risk of infection with COVID-19, more than 15 million undocumented immigrants and US citizens who lived in mixed-status households were ineligible to receive direct cash payment support under the 2020 CARES Act.[2] This exclusion from federal relief is part of the continuation of the systematic barring of undocumented immigrants since the passing of the Personal Responsibility and Work Opportunity Act of 1996 that denied or limited the eligibility of immigrants for federally funded programs.[3] Undocumented immigrants were eligible for no-cost COVID-19 testing, treatment, and vaccination.[2] Anti-immigrant policies and rhetoric, however, likely discouraged the use of health care services available to undocumented immigrants.[4]

STATE AND LOCAL GOVERNMENTS EXPAND HEALTH COVERAGE

Previous research has shown that documentation status is one of the main contributors to health and health care inequities and a key predictor of uninsured status for immigrants.[5,6] Lack of health insurance coverage among undocumented immigrants is associated with delays in seeking health care and underuse of cost-effective health care services.[7,8] With federal inaction regarding regularizing the stay of undocumented immigrants, several state and local governments have taken action to address the health care needs of undocumented immigrants. State and local policies are particularly salient because most undocumented immigrants live in a few states and metropolitan areas. For instance, approximately 63% of undocumented immigrants live in only six states (California, Texas, Florida, New York, New Jersey, and Illinois), and almost 82% of undocumented immigrants live in only 178 counties.[9] With declining federal support for immigrant health coverage, state and local safety net providers have had to assume the responsibility to offer health care and other basic public services to undocumented immigrants.

In this issue of *AJPH*, DeGarmo et al. (p. 1757) analyze state-level legislation targeting undocumented immigrants between November 2021 and August 2021. The authors used a systematic search method to identify and classify state bills that addressed the needs of undocumented immigrants during the

COVID-19 pandemic. Their main findings were that the legislatures of 13 states proposed a total of 66 bills classified under health-related services, job security and employment benefits, and monetary assistance. However, only 17 of these bills ultimately became law.

Although it is noteworthy that 94% of new legislation is protective of undocumented immigrants, this must be contrasted with a couple of policy dilemmas. First, federal aid, including the stimulus checks of 2020, generally excluded undocumented immigrants by default. Second, not all states that had large numbers of undocumented immigrants adopted protective legislation for them. For instance, California, New York, and Illinois were among the 13 states where protective legislation was proposed; however, states with large populations of undocumented immigrants, such as Texas or Florida, were among the 37 states maintaining a prepandemic status quo that left millions of undocumented immigrants at high risk.

STATE POLICIES FOR COVERAGE DURING PANDEMIC

This study is an important contribution to our understanding of how state policies aimed to address the increased vulnerabilities experienced by undocumented immigrants during the COVID-19 pandemic. Although the accounting of state legislation is an insightful and important metric, it has limitations. The approval of bills and resolutions to improve access to public resources for undocumented immigrants is an important first step; however, policies and programs need to be effectively implemented and evaluated to determine their effects. It remains unclear what the community health impacts were of these bills and resolutions that ultimately passed.

Moreover, providing health care to undocumented immigrants should be in the spirit of health as a human right and not simply a way of getting undocumented immigrants tested and vaccinated to prevent infectious diseases among citizens. As the authors point out, there was much conjecture among politicians and others about COVID-19 being spread by undocumented immigrants. Public health history in the United States has shown that medical professionals and others advocated for the health rights of Black people so that they would not spread disease to White people, as opposed to advocating for their health care as a human right.[10] We should not be repeating this history for immigrants or any other minoritized population.

Health policies and programs need to be improved so that immigrants trust health care providers and systems. Minoritized populations, including immigrants and especially undocumented immigrants, experience discrimination in health care.[11] Misinformation also reduces the reach and effective implementation of laws and policies targeting undocumented immigrants. For instance, a recent study estimated that 108 000 to 193 000 Latino immigrants without green cards in California did not enroll in Medicaid despite their eligibility, likely because of fear of the public charge rule even though the Biden administration reversed the change in its definition by the Trump administration.[4] Likewise, anti-immigrant rhetoric and polices likely contributed to lower COVID-19 testing, vaccination, and treatment uptake because of mistrust or fear of deportation of themselves or a friend or a family member, regardless of the state law.

EXPANDING HEALTH CARE IN STATES THAT LACK COVERAGE

One important finding of the study is that the 13 states that introduced bills and resolutions to help undocumented immigrants in the context of COVID-19 had large or rapidly growing undocumented immigrant populations. However, many states have not introduced any COVID-19–related public health legislation to help immigrants, and many of these states are conservative states where undocumented immigrants are working on the front line in essential jobs, especially in the food industry (e.g., farming, processing, distribution, retail). Although it is laudable that some states are starting to enact legislation to protect undocumented immigrants in terms of their health and social welfare, variation across states is widespread. It is also possible to find variation within states. For example, even though California has enacted progressive legislation to protect undocumented immigrants, variability remains in how laws and programs are implemented across counties, and implementation tends to vary by political party line.[7]

The importance of the study's findings is highlighted by two new threats. First, as the COVID-19 pandemic continues, another epidemic has emerged: the monkeypox virus. The disease carries high levels of stigma, requires up to 28 days of isolation, and has disproportionately affected Black and Latino populations, which underlies the need for stronger protections for undocumented immigrants. Simultaneously, the Biden administration's attempt to limit the number of individuals targeted by immigrant enforcement agencies has been curtailed by ongoing litigation, which raises the possibility that immigration arrests may resurge to prepandemic levels. Likewise, the sunsetting of Title 42, a public health law that has been used more for immigration enforcement than for COVID-19 prevention, has also been delayed by the actions of state and national officials who came to the program's defense. This will likely have a chilling effect on screening and contact tracing, which is similar to the chilling effects produced by changes to the public charge rule in 2019 by the Trump administration.[4] In fact, one particularly worrisome issue is that few state-level protections attempted to prohibit immigration authorities from accessing contact tracing data directly.

It is critical that US public health policy be proactive in the face of these public health crises, rather than being reactionary after new diseases have taken hold. The study underscores the need for a universalist approach to legal protections. Our society's public health fabric is only as strong as its weakest link, and, when we exclude groups a priori, we facilitate the resurgence of disease. The legislation profiled in this study should serve as a blueprint for other governments seeking to navigate the landscape of immigration policy and laws during the age of climate change and increasing infectious disease pandemics. AJPH

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The Harmful Impacts of Anti-Immigrant Policies on Maternal and Child Health

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See also Wang et al., p. 1747.

In January 2017, a draft executive order leaked that aimed to limit an immigrant's ability to gain lawful permanent residence status if they used public benefits, including Medicaid and the Supplemental Nutrition Assistance Program (SNAP). This policy was then included in a proposed regulation issued in 2019 that was implemented in February 2020. In a new study by Wang et al. (p. 1747 in this issue of *AJPH*), the leak of the draft executive order was found to be associated with delayed Medicaid enrollment and adverse maternal and child health outcomes in New York State. Sadly, these are not isolated findings, as these results align with previous research on the multifaceted challenges immigrants experience in accessing health care in the United States.[1]

High-quality health care is important for optimal maternal and child health outcomes, particularly throughout the stages of pregnancy (i.e., preconception, prenatal, and postpartum).[2] Disparities among immigrant women in access to pregnancy-related services have been well documented: immigrant women are less likely to have a usual source of care and are more likely to have inadequate and delayed initiation of prenatal care[3] than are US-born women. Immigrant mothers to-be encounter structural inequities, including language and cultural barriers, adverse or unequal treatment, financial burdens, and anti-immigrant policies, that are collectively associated with adverse birth outcomes.[4] The Wang et al. study analyzed one potential policy change—the public charge rule.

PUBLIC CHARGE

As described by Wang et al., "public charge" was largely undefined in US immigration law since its implementation in 1882. It was not until 1999 that federal regulatory guidance provided a limited definition of public charge related to those who depended on federal benefit programs for their income or required long-term institutionalized care. In January 2017, the presidential administration of Donald Trump proposed changing the definition of public charge simultaneously with other broad federal anti-immigration policies, such as accepting reduced numbers of refugees and banning noncitizens from several predominantly Muslim countries from entering the United States.

Although immigration policy is primarily a federal issue, states may further develop and implement policies that can either include or exclude immigrants.[5] Exclusionary policies, at all levels of government, can contribute to systemic racism, and enforcement of these exclusionary policies has been found to have detrimental effects on immigrants and their families—as well as on US citizens, particularly those in mixed status households.[6] For instance, US Immigration and Customs Enforcement raids have been found to be associated with a greater risk of preterm low birthweight among both US-born and immigrant Latina mothers.[7] Furthermore, separate from specific policy initiatives, the 2016 US presidential election was found to be associated with an increase in preterm births among US Latina women,[8] foreign-born Latina women (specifically with Mexican or Central American ancestry), and women from the Middle East and North Africa.[9]

In this context, Wang et al. found that, after the memo was leaked, noncitizen pregnant mothers were more likely than were citizen mothers to delay prenatal enrollment in Medicaid, and their infants were more likely to have lower birth weights than were infants of citizen mothers. Of note, these changes occurred in New York, a state that has historically had more inclusive health and welfare immigrant policies than have other states,[5] suggesting that the adverse outcomes detected by Wang et al. may have been even worse in other states.

STRENGTHS AND LIMITATIONS

The strengths of the study include the authors' use of detailed data on health care enrollment and utilization and

AJPH December 2022, Vol 112, No. 12

health outcomes from a large Medicaid program and multiple sensitivity analyses to probe the robustness of the results. It is not entirely clear whether the extent of prenatal coverage and care changes seen in the study were large enough to explain the observed changes in birth weight, although as the authors' note, research shows that psychosocial stressors themselves—such as a hostile policy environment—can be a contributing factor to adverse pregnancy outcomes.[10]

A few limitations of this study include missing data on citizenship status among many enrollees (as high as 30% in 2019), which appeared to increase over time and could have confounded the findings, even after imputation for missing values. In addition, the fact that changes in outcomes for noncitizens began to appear even before the 2016 election raises some question about secular trends and the causal role of the January 2017 leak; however, previous analyses have suggested that anti-immigrant rhetoric during the 2016 presidential election campaign itself may have changed health care utilization and health outcomes for immigrants. Lastly, some of the findings on health outcomes depend on the model presented. For instance, the findings differ based on imputed versus nonimputed maternal citizenship status and New York City versus non–New York City enrollees. However, despite these limitations, the differential changes that Wang et al. observed between noncitizens and citizens in most models is highly suggestive of a link to the policy in question.

FEDERAL ACTION ADDRESSING DISPARITIES

Because immigrant health is shaped by the context of immigration policies, inclusive and protective policies for immigrants are important tools that may improve health equity. On September 8, 2022, the US Department of Homeland Security issued a final rule on new public charge regulations that would largely codify 1999 field guidance governing public charge determinations, but with some changes.[11] This new rule allows eligible individuals to enroll, without harmful immigration consequences, in programs such as Medicaid (except for long-term institutionalization at government expense), the Children's Health Insurance Program, and SNAP.

Furthermore, immigrant-focused policies exist in a broader framework of policies and social determinants that affect the health and well-being of immigrant communities, which is evident in the current presidential administration's approach to a range of issues designed to increase health equity. President Joseph Biden signed Executive Order 14009, "Strengthening Medicaid and the Affordable Care Act" (January 28, 2021), and Executive Order 13985, "Advancing Racial Equity and Support for Underserved Communities Through the Federal Government" (January 20, 2021), as part of a broader effort to address coverage gaps and structural inequities that disproportionately affect immigrants and other communities. This effort includes the implementation of the American Rescue Plan Act (Pub L No. 117–2; March 11, 2021) provision that enables states to provide continuous Medicaid eligibility for 12 months after pregnancy and a comprehensive approach to addressing social determinants of health.[12] In terms of access to coverage, although detailed information on insurance rates among immigrants is not yet available, national survey data in early 2022 showed that the US uninsured rate had reached its lowest level ever, indicating that there are now more than 5 million more US residents with health coverage than there were in 2020.[13]

As suggested by the study of Wang et al. and the wide-ranging related literature that preceded it, anti-immigrant rhetoric and policies are detrimental to society, as they contribute to increased psychosocial stress, lower access to care, and negative health effects, including adverse maternal and child health outcomes. Efforts by the current presidential administration are a step toward improving maternal and child health outcomes among immigrants residing in the United States. More broadly, these policies can help dismantle persistent health disparities, including those affecting the more than 40 million immigrants living in the United States. AJPH

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Immigrant-Inclusive Policies Promote Child and Family Health

Allison Bovell-Ammon, MDiv, Stephanie Ettinger de Cuba, PhD, MPH, and Diana B. Cutts, MD

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See also Miller et al., p. 1738.

In this issue of *AJPH*, findings by Miller et al. (p. 1738) suggest that anti-immigrant rhetoric and proposed changes to public charge during the early years of the Trump administration significantly reduced federal nutrition assistance program participation among mixed-status immigrant households. Mixed-status households in states with the most, compared with the least, generous eligibility provisions for noncitizens had greater declines in Supplemental Nutrition Assistance Program (SNAP) participation, and those in moderately generous states saw declines across SNAP and school meals programs. Public charge is a determination made when some potentially eligible noncitizens seek legal permanent residency. Trump-era changes to public charge included an expansion of assistance programs considered indicative of whether the applicant is deemed likely to be dependent on long-term government assistance in the future (i.e., a "public charge"). Prior to final rule issuance in August 2019, the proposed rule and several leaked drafts, as well as speculation about the scope, time frame, and contents of the rule change, perpetuated fear in immigrant communities. Miller et al.'s results indicating significant participation reductions in SNAP, the National School Lunch Program (NSLP), and the School Breakfast Program (SBP) are concerning, given robust evidence demonstrating these programs' health and educational benefits. These findings suggest potentially harmful long-term consequences of anti-immigrant rhetoric and regulatory changes, underscoring the urgent necessity of implementing policy solutions that promote equitable assistance program access without fear.

HEALTH EFFECTS OF REDUCED BENEFIT PARTICIPATION

Decades of research show that SNAP participation is associated with health benefits across the life span, including positive birth outcomes, healthy cognitive development among children, and good overall health status and reduced acute health care use and spending for children and adults, in addition to reducing food insecurity.[1–3] Beneficial health impacts of SNAP participation in childhood persist into adulthood.[2] School meal programs are associated with positive health and education outcomes among children. NSLP is linked to reduced rates of poor health and obesity among school-age children and improved attendance, behavior, and academic achievement.[4] SBP is associated with improved nutrient intake, better student mental health, and positive education outcomes.[5] Given these public health considerations, paired with the fact that more than one-fourth of children in the United States have at least one immigrant parent, maintaining consistent access to federal nutrition assistance programs is essential for promoting optimal population health.

Although Miller et al. did not find changes in food security, other research demonstrates increased rates of food insecurity among families with immigrant mothers following the 2016 election.[6] Both the final expanded rule, which took effect in fall 2019, and the COVID-19 pandemic occurred after the study period presented in Miller et al.'s article[7]; still, following these events, chilling effects in federal assistance program participation persist. Given rising economic hardships resulting from the COVID-19 pandemic, Miller et al.'s findings become only more relevant for ensuring that families with noncitizens continue to be able to afford basic needs. Lessons from the pandemic response may further illuminate necessary action. Noncitizen and mixed-status families have faced an increased risk of COVID-19–related poor health outcomes and economic hardships during the pandemic compared with US-born households while being less likely to benefit from COVID-19–related protections and relief policies.[8,9]

SYSTEMIC, POLICY CHANGES NEEDED

The Biden administration has taken steps to reverse harmful changes to

AJPH December 2022, Vol 112, No. 12

the public charge rule and has finalized a rule returning the public charge definition to the 1999 precedent, which narrowly focused on specific cash benefits and public long-term institutionalization and excluded other housing, food, and health care programs (Miller et al.). This effort was undertaken by the administration to stem well-documented chilling effects in health and assistance programs among immigrants and their families. Reversal is an important step toward alleviating chilling effects, but issuance of the new public charge regulation alone is unlikely to ameliorate harms inflicted upon immigrant communities across decades of US policy.

Miller et al. rightfully emphasize effects on public assistance participation among noncitizens following exclusionary policymaking efforts in the late 1990s and the ways state-level responses interacted with federal level changes in families' lives. In addition to existing public assistance program eligibility restrictions and changes to public charge, increasing efforts across the nation to criminalize immigrant communities, separate families, and marginalize immigrants through xenophobic rhetoric have resulted in significant harm that is not easily undone.[11] Responding to the public health issue of xenophobia and anti-immigrant policymaking will require a robust response across all levels of government and society.[11]

Policy and programmatic solutions responsive to the needs and requests of immigrants themselves are important for advancing equity and immigrant inclusion. In addition to comprehensive immigration reform that creates a path to citizenship, eliminates family separation, and lifts pandemic-era border restrictions on asylum seekers, federal legislation that simplifies eligibility, is inclusive, and eliminates barriers to assistance programs is paramount. The complex patchwork of eligibility rules across public assistance programs creates significant confusion—not just for immigrant families in need of support but also for public assistance workers, service organizations, and legal professionals, not to mention the general public. Removing all immigration-related rules from eligibility determinations would provide the most seamless and health-promoting access to the essential support provided by SNAP, school meals, and other public assistance programs. Experience gained during the pandemic shows implementation of universal school and child care meals nationwide would mean that all children, regardless of immigration status, have access to healthy meals without unnecessary and costly administrative burden. For SNAP, important progress toward more inclusive policy would include lifting the five-year bar that prevents otherwise eligible, lawfully present noncitizens who have resided in the United States for less than five years from accessing SNAP and other health-promoting federal programs. These changes are critically important investments in the current and future health of children in the United States.

Congress and the current administration have several imminent opportunities for enacting transformative policy improvements. These include current efforts to develop a national strategy to end hunger by 2030 as part of the White House Conference on Hunger, Nutrition, and Health; ongoing Child Nutrition Act reauthorization deliberations; and the forthcoming farm bill debate. Intentionally focusing on addressing the marginalization of immigrant families, including specific attention to both mixed-status and noncitizen families, in federal policy discussions is critical to reverse harms documented by Miller et al. and many others. In addition to federal policy change, investment in and support for community-based groups with a track record of responding to the needs of noncitizen and mixed-status families is important for further bolstering immigrant health and the health of the more than one fourth of children in the United States with immigrant parents. Miller et al. hypothesize that these community groups may have been key in promoting food security among mixed-status families despite declines in federal assistance program participation.

Finally, although these changes and investments are important, an adequate response to generations that have experienced historical bias and trauma requires further action. Rebuilding trust in public institutions and reversing adverse outcomes will require sincere engagement with trusted immigrant-led community groups and elevation of a diversity of immigrant voices in decision-making. Following the leadership of immigrants is essential not only for establishing trust but also for ultimately ensuring that equitable policies are enacted, evaluated, and continuously improved. Only then can we achieve truly equitable child and family health for all families in the United States. AJPH

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CONFLICTS OF INTEREST

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The 2016 Presidential Election, the Public Charge Rule, and Food and Nutrition Assistance Among Immigrant Households

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See also Bovell-Ammon et al., p. 1735.

Objectives. To investigate whether the 2016 US presidential election and the subsequent leak of a proposed change to the public charge rule reduced immigrant families' participation in food and nutrition assistance programs.

Methods. We used nationally representative data on n = 57 808 households in the United States from the 2015–2018 Current Population Survey–Food Security Supplement. We implemented difference-in-difference-in-difference analyses to investigate whether the election and proposed rule change produced decreases in immigrant families' participation in food and nutrition assistance programs and whether such decreases varied according to state policy generosity toward immigrants.

Results. Findings indicate significant and large decreases in Supplemental Nutrition Assistance Program, School Breakfast Program, and National School Lunch Program participation among immigrants in moderately generous states but no changes to receipt of food assistance from nongovernmental sources or to household food insecurity.

Conclusions. Both anti-immigrant rhetoric and the perceived threat of policy enactment can be enough to produce chilling effects that have potentially serious implications for the health of immigrant households and thus the health of the nation. (*Am J Public Health*. 2022;112(12):1738–1746. https://doi.org/10.2105/AJPH.2022.307011)

AJPH December 2022, Vol 112, No. 12

Shortly after the 2016 US presidential election and following a campaign by Donald Trump characterized by a decidedly hostile tone toward immigrants and their families,[1,2] a draft of a Trump administration executive order was leaked that proposed changes to the public charge rule. For immigrants applying for legal permanent residence, this change would have greatly expanded the number of public assistance programs for which previous receipt of benefits could be counted in determining whether they were likely to become a future public charge, a designation that could lead to a rejection of their applications.[3] In addition to federal cash assistance and public long-term care, which had long been used in the public charge determination, the 2016 proposed change would have included use of programs such as the Supplemental Nutrition Assistance Program (SNAP), Medicaid, and federal housing assistance (although not federal school meals programs).[4] Early well-publicized drafts of the rule change also suggested that the participation of family members such as US-born children would also be newly used in the public charge determination.[5]

Based in part on decreases in participation in public programs that followed the 1996 Personal Responsibility Work Opportunity Reconciliation Act (PRWORA; Pub L No. 104–193),[6] which eliminated eligibility for public assistance for most legally resident immigrants,[7] both the 2016 election and the proposed rule change generated renewed concern about "chilling effects."[8] In a legal context, this term typically describes "undesirable discouraging effects or influences." Here we use the term to mean immigrants foregoing public benefits to which they were legally entitled. Indeed, complementing media coverage, researchers found that the leak of the proposed rule changes was associated with sizable decreases in SNAP participation among recent immigrant families with younger children[9] and Medicaid participation in counties with larger noncitizen populations. When a modified version of the public charge rule change was eventually implemented in December 2018, researchers found that 20% of low-income immigrant adults reported avoiding a public benefit program because of perceived threats to their residence status.[10] There was also evidence of large-scale avoidance of SNAP and Medicaid by immigrant essential workers.[11]

Unlike in 1996, there were no immediate changes to eligibility for public benefits in the early days of the Trump presidency. Rather, the leaked draft executive order outlined changes to the public charge rule that would create potentially serious consequences for the receipt of federal public assistance. This, coupled with increasingly harsh rhetoric and other executive orders that targeted immigrants,[12] led to renewed fear of decreases in program participation. In addition, misinformation and confusion propagated in part by news media appeared to have added to hesitation about participating in public assistance.[13,14]

Nonetheless, an important insight from research on PRWORA is that the experience of chilling effects is likely to vary by the composition of immigrant households.[15] For instance, studies reported that there were pronounced decreases in program participation among mixed status households (those with citizen children and noncitizen adults),[16,17] though other research indicated that these decreases may have been because of changing food stamp benefits rates[15] and changes to naturalization.[7] PRWORA era research also signals the importance of state policies to the potential for chilling effects. In the late 1990s, some states provided benefits to immigrants in response to their loss of eligibility for federal programs, which lead to reductions in program participation.[18,19]

Building on recent evidence[9–11,14,20] and this previous research, we provide a definitive assessment of the effects of the 2016 election and the leak of the proposed public charge rule change on immigrant families' food insecurity and federal food and nutrition assistance use. To our knowledge, our study is the first to do so using nationally representative data on US households. We consider the effects of the 2016 election and the rule change leak on mixed status households and whether any effects vary by states' generosity in providing benefits to immigrant households.

As with previous research,[7,16,17] we expected to see the strongest chilling effects in mixed status households (i.e., those with noncitizen parents and citizen children) because they might especially fear the serious disruptions an adverse public charge determination would cause. While actual changes in eligibility may have driven behavior after PRWORA, we investigated instead whether an increased climate of anti-immigrant sentiment and a proposed change to policy suppressed participation. Furthermore, we hypothesized that states' generosity toward immigrants in 2016 might have actually encouraged a retreat from federal benefits if immigrant households believed they could switch to a state program in lieu of a federal one.

METHODS

We used data from the Current Population Survey (CPS)–Food Security Supplement (FSS). Each month, the CPS is administered to a national sample of households, which are representative of the noninstitutionalized US population. The FSS is administered each December and contains detailed data on household food expenditures and the use of both governmental and nongovernmental food assistance.

Using the Integrated Public-Use Microdata Series,[21] we constructed a preliminary analytic sample of n = 150 853 households using data from the 2015 to 2018 waves of the CPS–FFS, a period including the 2 years before the 2016 election (2015–2016) and the first 2 years of Trump's presidency (2017–2018). To focus on those most likely to take advantage of governmental programs and nongovernmental aid, we dropped n = 91 213 families with incomes greater than $40 000 per year. Finally, we dropped n = 1810 households in which no members were citizens. Our final analytic sample had n = 57 808 households and subsamples of n = 10 832 and n = 10 811 households with school-aged children (aged 5–17 years) in our respective analyses of the National School Lunch Program (NSLP) and the School Breakfast Program (SBP).

Measures

Outcomes. We coded variables indicating participation in multiple federal food and nutrition assistance programs. First, we created a dichotomous measure of participation in SNAP, the largest of the US Department of Agriculture's (USDA's) food and nutrition assistance programs,[22] coded as 1 for households who had received SNAP benefits since December of the previous calendar year and 0 otherwise. Next, for households with school-aged children, we created additional dichotomous indicators for whether respondents reported that children in the household received free or reduced-price meals from the NSLP or SBP in the past month. We coded receipt of food assistance from nongovernmental sources as 1 if respondents reported that anyone in the household had gotten emergency food from a church, food pantry, or food bank or had eaten at a soup kitchen in the past month. Finally, and based on the 18-item Food Security Module, which is included in the CPS–FFS, we used USDA guidelines[23] to create a 0–1 indicator for household food insecurity over the previous 12 months. We provide full information about the construction of these and other key variables in Appendix A (available as a supplement to the online version of this article at http://www.ajph.org). Table 1 provides descriptive information on all study variables.

Household citizenship status. We assigned CPS–FFS households to 1 of 3 categories: all-citizen, noncitizen, and mixed status households, in which some members were citizens and others were not. However, preliminary analyses showed divergent preelection trends in our outcomes of interest between noncitizen households and the 2 other groups, indicating a violation of a key assumption undergirding our analytic approach.[24] For this reason, we elected to drop noncitizen households from our analyses.

State generosity. Based on previous research,[15,19,25] we measured the

TABLE 1— Descriptive Statistics for the Analytic Sample (n = 57 808): United States, 2015–2018 Current Population Survey–Food Security Supplement

	% (No.) or Mean ±SD	Range
SNAP	20.3 (11 735)	0–1
Nongovernmental food	11.0 (6 359)	0–1
NSLP (n = 10 832)	59.7 (6 471)	0–1
SBP (n = 10 811)	51.7 (5 587)	0–1
Food insecurity	22.4 (12 949)	0–1
Mixed status household	6.2 (3 584)	0–1
State policy generosity		
Least	26.8 (15 493)	0–1
Moderate	58.9 (34 049)	0–1
Most	14.4 (8 324)	0–1
Respondent race/ethnicity		
Non-Hispanic White	68.1 (39 367)	0–1
Non-Hispanic Black	11.4 (8 093)	0–1
Non-Hispanic American Indian/ Alaska Native	1.6 (925)	0–1
Non-Hispanic Asian	2.4 (1 387)	0–1
Non-Hispanic Hawaiian/Pacific Islander	0.2 (116)	0–1
Non-Hispanic other race	1.5 (867)	0–1
Hispanic any race	12.1 (6 995)	0–1
Respondent in labor force	44.2 (25 551)	0–1
Respondent marital status		
Married, spouse present	28.1 (16 244)	0–1
Married, spouse absent	1.9 (1 098)	0–1
Separated	3.7 (2 139)	0–1
Widowed	21.7 (12 544)	0–1
Divorced	18.0 (10 405)	0–1
Never married	26.6 (15 377)	0–1
Respondent education		
< High school	17.1 (9 885)	0–1
High school	36.7 (21 216)	0–1
Some college	20.8 (12 024)	0–1
Associate's degree	10.3 (5 954)	0–1
Bachelor's degree or more	15.1 (8 729)	0–1
Household size	2.063 ±1.345	1–14
Respondent age	55.08 ±18.09	15–85
Family income (in 2020 US$)	23 518.5 ±11 463.8	2 589.5–41 258.2
Family income < 185% federal poverty threshold[a]	65.0 (37 575)	0–1
State policy index (lagged 1 y)	0.657 ±0.851	−0.571–2.882

Note. NSLP = National School Lunch Program; SBP = School Breakfast Program; SNAP = Supplemental Nutrition Assistance Program.

[a]Federal thresholds defined by the US Census Bureau for 2015–2018.

number of assistance programs (0–3) that states had established for immigrants as of 2017. Specifically, we measured whether immigrants were eligible for (1) state food and nutrition assistance programs (n = 6 states in 2017), (2) state replacement for the federal Supplemental Security Income program (n = 5 states), and (3) state replacement for the federal Temporary Assistance for Needy Families program (n = 22 states). In addition, we coded whether states had chosen to take up the federal option to expand Medicaid and Children's Health Insurance Program coverage to immigrant families who had been in the country for fewer than 5 years (n = 32 states). We coded states as less generous if they had not adopted any of these policies (n = 14), as moderately generous if they had adopted 1 or 2 policies (n = 29), or as most generous if they had adopted 3 or 4 of these policies (n = 7).

Covariates. In all analyses, we controlled for potential confounders, including respondent race/ethnicity, labor force participation, marital status, education level, household size, age, family income, and an indicator for whether household income was below 185% of the US Census Bureau's poverty thresholds in the appropriate survey year (2015–2018). We also included a standardized index ($\alpha = 0.821$; mean = 0; SD = 1) of state-based controls using data from the University of Kentucky Center for Poverty Research National Welfare Database.[26] We lagged all measures by 1 year before including them in the index.

Statistical Analysis

We used difference-in-difference-in-difference (DDD) analyses. Difference-in-differences (DD) approach is a commonly adopted quasiexperimental method used to generate causal estimates of policy changes or other interventions. The central insight of the approach was that we could detect chilling effects by comparing changes in program participation rates for mixed status households before and after the 2016 election (the first difference) while accounting for whatever secular changes occurred in the outcome over the same period among citizen households (the second difference), whose program participation was unlikely to be affected by the election or proposed change to the public charge rule. In our analyses, we extended this basic DD approach by examining whether effects were more or less pronounced among immigrant households living in states with policies that were more generous to immigrants. In these models, our DDD estimates were the difference between the DD for mixed status families in moderate- and high-generosity states and the DD for mixed status families in low-generosity states. These analyses allowed us to investigate potential chilling effects after accounting for secular trends among citizen households and among mixed status households in the lowest generosity states, whose participation in public programs may have been unaffected by the election and proposed rule change.

We implemented our DDD approach using linear regressions[24] that included 3-way interactions between time (0 = 2015/16, 1 = 2017/18), the indicator for household mixed status, and state policy generosity (i.e., less, moderate, most). For all analyses, we included controls for the variables described in the Covariates section, clustered our SEs at the state level, and used probability weights supplied in the CPS–FFS to generate nationally representative estimates. We examined outcome trends before 2016 and used event study analysis to test the parallel trends assumption for each of our outcomes. We also conducted a series of sensitivity analyses, rerunning our analyses using probit models to assess whether our results varied depending on functional form, and again after including state and year fixed effects as a further check against bias from endogeneity. We completed all analyses using Stata version 16 (StataCorp LP, College Station, TX).

RESULTS

Unweighted descriptive statistics are shown in Table 1. Over the study period, 20.3% of all sample households had received SNAP benefits in the previous calendar year, 11.0% had received some type of nongovernmental food assistance, and 22.4% were food insecure over the previous year. More than half of households with school-aged children reported participation in the NSLP (59.7%) and the SBP (51.7%).

Results from our parallel trends and event study analyses in Appendix B (available as a supplement to the online version of this article at http://www.ajph.org) do not reveal any meaningfully different pre-2016 group trends for any of our outcomes. Weighted results from our DDD models with our analytic sample of CPS–FFS households are presented in Table 2. The table shows parameter estimates for our primary study variables and their interactions. The primary results of interest are the DDD estimates, which we show in the final rows of the table. Full regression results for all models are available on request.

Table 2 shows that the 2016 election and leak of the proposed rule change produced decreases in SNAP participation

TABLE 2— Effects of the 2016 Presidential Election and Leak of the Proposed Public Charge Rule Change: United States, 2015–2018 Current Population Survey–Food Security Supplement

	SNAP (n = 57 808), b (95% CI)	Nongovernmental Food Aid (n = 57 808), b (95% CI)	NSLP (n = 11 332), b (95% CI)	SBP (n = 11 309), b (95% CI)	Past-Year Food Insecurity (n = 57 808), b (95% CI)
Postelection	−0.023 (−0.045, −0.002)	−0.019 (−0.035, −0.004)	0.005 (−0.025, 0.034)	0.0023 (−0.023, 0.068)	−0.025 (−0.038, −0.012)
Mixed status Household	−0.112 (−0.159, −0.065)	−0.025 (−0.060, 0.010)	−0.048 (−0.132, 0.036)	−0.055 (−0.195, 0.085)	−0.029 (−0.070, 0.011)
Least generous (Ref)	0	0	0	0	0
Moderate generosity	0.018 (−0.005, 0.042)	−0.013 (−0.039, 0.013)	−0.010 (−0.049, 0.029)	−0.009 (−0.059, 0.041)	−0.017 (−0.035, 0.002)
Most generous	0.037 (0.003, 0.071)	0.021 (−0.006, 0.048)	−0.021 (−0.064, 0.021)	−0.061 (−0.128, 0.005)	−0.017 (−0.035, 0.001)
Postelection × mixed status	0.071 (0.029, 0.113)	0.013 (−0.030, 0.056)	0.099 (0.007, 0.192)	0.095 (−0.045, 0.234)	0.001 (−0.065, 0.067)
Postelection × least generous (Ref)	0	0	0	0	0
Postelection × moderate generosity	0.022 (−0.001, 0.044)	0.022 (0.005, 0.039)	0.024 (−0.016, 0.064)	0.012 (−0.044, 0.067)	0.021 (0.003, 0.038)
Postelection × most generous	−0.010 (−0.039, 0.019)	0.004 (−0.012, 0.019)	−0.022 (−0.075, 0.030)	0.019 (−0.031, 0.069)	0.013 (−0.008, 0.034)
Mixed status × least generous (Ref)	0	0	0	0	0
Mixed status × moderate generosity	0.039 (−0.012, 0.089)	−0.017 (−0.056, 0.022)	0.059 (−0.036, 0.154)	0.069 (−0.078, 0.217)	−0.028 (−0.076, 0.019)
Mixed status × most generous	0.046 (−0.019, 0.111)	−0.023 (−0.057, 0.011)	0.135 (0.042, 0.229)	0.136 (−0.019, 0.291)	0.012 (−0.046, 0.071)
Postelection × mixed status × least generous (Ref)	0	0	0	0	0
Postelection × mixed status × moderate generosity	−0.073 (−0.128, −0.019)	−0.022 (−0.073, 0.028)	−0.126 (−0.234, −0.019)	−0.160 (−0.311, −0.008)	−0.001 (−0.076, 0.074)
Postelection × mixed status × most generous	−0.068 (−0.129, −0.006)	0.011 (−0.038, 0.060)	−0.059 (−0.165, 0.047)	−0.092 (−0.239, 0.055)	−0.039 (−0.125, 0.047)

Note. CI = confidence interval; NSLP = National School Lunch Program; SBP = School Breakfast Program; SNAP = Supplemental Nutrition Assistance Program.

among mixed status households in states with moderately or most generous policies, as hypothesized. The predicted size of these decreases was quite large, 7.3 and 6.8 percentage points, respectively. Similarly, DDD estimates indicated decreases in NSLP participation of 12.6 percentage points and SBP participation of 16.0 percentage points among mixed status households in moderate generosity states. Parameter estimates for NSLP and SBP participation for mixed status households living in the most generous states were negative but not statistically significant.

Notably, despite decreases in participation in 3 national nutrition programs, the 2016 election and leak of the proposed rule change did not result in significant changes to household food insecurity for mixed status households. To assess whether the lack of significant findings was related to our definition of household food insecurity and taking advantage of the 10 adult-referenced and 8 child-referenced questions in the USDA Food Security Module, we reran our models using past-month and past-year household, adult, and child food insecurity (results available on request). Across all of these models, we found no evidence that the 2016 election or the leak of the proposed rule change had any significant impact on food insecurity. Likewise, we found no evidence of chilling effects for receipt of nongovernmental food aid.

Table 3 presents the results of our sensitivity analyses. For interpretability, the table presents only DDD parameter estimates. For each outcome, the first column presents again the results from our main analyses. Across outcomes, the results shown in the table indicate that our main results are not sensitive to assumptions about functional form and are not biased because of unobserved characteristics of states or years of measurement. In fact, Table 3 indicates strong consistency of both pattern and magnitude of parameter estimates. The sole exception is minor: the parameter estimate for SNAP participation in the most generous states from the probit model, which just misses the cutoff for statistical significance ($P = .054$).

DISCUSSION

To our knowledge, this study is the first to use nationally representative data to investigate whether the 2016 presidential election and subsequent leak of a proposed change to the public charge rule resulted in chilling effects in immigrant households' participation in food and nutrition programs. Building on intuition developed in earlier, PRWORA era research, we pooled data from 2 years before and 2 years after the election and used DDD models to assess whether the election and proposed rule change produced changes in household food insecurity and in the receipt of SNAP, school meal programs, and nongovernmental food aid that varied by state policy generosity.

Similar to previous work,[11,14,20] our most consistent findings are for mixed status households living in states that had adopted a moderately generous set of policies toward immigrants. For this group, we found that the combination of the 2016 election and the proposed rule change produced sizable decreases in SNAP participation (–7.3 percentage points), NSLP participation (–12.6 percentage points), and SBP participation (–16.0 percentage points). Compared to participation rates in SNAP (20.3%), NSLP (59.7%), and SBP (51.7%), participation rates in our sample of low-income households, these estimates represent substantial and serious decreases in participation in 3 of the primary federal programs to fight food insecurity among households with children. It is surprising, then, that our analyses did not find any change in household food insecurity for mixed status households in these states. One explanation might be an increased propensity for immigrant households to receive food assistance from nongovernmental sources. However, our analysis found no change in receipt of food from nongovernmental sources such as churches, food banks, food pantries, or shelters. A further explanation is that mixed status households turned to informal social supports to help meet food needs and thus were able to stave off increases in food insecurity. Unfortunately, the FSS does not collect information on these types of supports, and so we could explicitly test for this possibility.

Even if immigrants turned to such supports, it is unlikely this aid would be consistent enough over time to completely prevent food insecurity if decreases in participation are sustained over time. Furthermore, even if eventual impacts on food insecurity are not realized, decreases in participation in SNAP and the 2 school meal programs are highly concerning in light of a growing body of research finding additional benefits to participation in these programs.[27–31] Complementing other research on the 2016 election,[9–11,20] our findings point to serious and ongoing negative impacts on public health related to anti-immigrant rhetoric and policy proposals that threaten the security of immigrant households.

Unexpectedly, we found little evidence of chilling effects for mixed status households in the most generous states, where we might have expected

TABLE 3— Effects of the 2016 Presidential Election and Proposed Rule Change: United States, 2015–2018 Current Population Survey–Food Security Supplement

	Main Model, b (95% CI)	Probit Model, b (95% CI)	State Fixed Effects, b (95% CI)	State and Year Fixed Effects, b (95% CI)
SNAP				
Postelection × mixed status × least generous (Ref)	0	0	0	0
Postelection × mixed status × moderate generosity	−0.073 (−0.128, −0.019)	−0.296 (−0.503, −0.089)	−0.081 (−0.133, −0.030)	−0.081 (−0.133, −0.030)
Postelection × mixed status × most generous	−0.068 (−0.129, −0.006)	−0.246 (−0.496, 0.005)	−0.071 (−0.128, −0.015)	−0.071 (−0.128, −0.014)
Nongovernmental food aid				
Postelection × mixed status × least generous (Ref)	0	0	0	0
Postelection × mixed status × moderate generosity	−0.022 (−0.073, 0.028)	−0.108 (−0.395, 0.178)	−0.029 (−0.081, 0.024)	−0.029 (−0.081, 0.024)
Postelection × mixed status × most generous	0.011 (−0.038, 0.060)	0.069 (−0.187, 0.326)	0.010 (−0.041, 0.062)	0.010 (−0.041, 0.062)
NSLP				
Postelection × mixed status × least generous (Ref)	0	0	0	0
Postelection × mixed status × moderate generosity	−0.126 (−0.234, −0.019)	−0.389 (−0.697, −0.080)	−0.128 (−0.236, −0.020)	−0.128 (−0.237, −0.020)
Postelection × mixed status × most generous	−0.059 (−0.165, 0.047)	−0.170 (−0.491, 0.150)	−0.053 (−0.157, 0.052)	−0.054 (−0.159. 0.051)
SBP				
Postelection × mixed status × least generous (Ref)	0	0	0	0
Postelection × mixed status × moderate generosity	−0.160 (−0.311, −0.008)	−0.455 (−0.864, −0.047)	−0.164 (−0.315, −0.013)	−0.164 (0.316, −0.013)
Postelection × mixed status × most generous	−0.092 (−0.239, 0.055)	−0.274 (−0.687, 0.138)	−0.089 (−0.234, 0.055)	−0.090 (−0.236, 0.055)
Past-year food insecurity				
Postelection × mixed status × least generous (Ref)	0	0	0	0
Postelection × mixed status × moderate generosity	−0.001 (−0.076, 0.074)	−0.012 (−0.270, 0.245)	−0.006 (−0.081, 0.069)	−0.006 (−0.081, 0.069)
Postelection × mixed status × most generous	−0.039 (−0.125, 0.047)	−0.122 (−0.413, 0.169)	−0.039 (−0.126, 0.048)	−0.040 (−0.126, 0.047)

Note. CI = confidence interval; NSLP = National School Lunch Program; SBP = School Breakfast Program; SNAP = Supplemental Nutrition Assistance Program.

reductions in participation to be greatest. The only evidence was a significant decrease in SNAP participation of 6.8 percentage points, although post hoc analysis indicated that this effect was not significantly different from the decrease for mixed status households in moderately generous states. Similar post hoc tests indicate that—although not significantly different from zero—the predicted decreases in NSLP and SBP participation for the most generous states were also not significantly different from those for moderately generous states. Although derived from previous work,[18,19] it may thus be that our system for classifying state generosity did not meaningfully distinguish between moderately and most generous states. Indeed, when we replicated our analyses by collapsing the moderately and most generous categories into 1 group, the pattern of results (available on request) was largely consistent. Thus, an important implication of this study is the need for policy researchers to continue to explore how the effects of national policy changes (or threats of policy change) interact with state-level policies and behaviors to affect health outcomes.

Limitations

Our study's results must be interpreted in the context of its limitations. Although we implemented a quasiexperimental approach that can control for unobserved heterogeneity, we relied on observational data and thus cannot definitively rule out potential bias. Furthermore, the limitations of survey data for analyzing program participation are well recognized. For this study, a particular additional challenge is the possibility that chilling effects are also realized in immigrant households' responses to survey questions. That is, immigrants fearing surveillance may have been less likely to report participation in government programs even if their actual behavior did not change. Although we do not consider this possibility very likely, both of these limitations underscore the importance of using administrative data on program participation to replicate the analyses and findings reported here.

Furthermore, we are unaware of any other comparable national data source that contains detailed information on our key study variables that does not rely on survey data. Finally, although we believe that our study design adequately captures the joint effects of the 2016 election and leaked proposed public charge rule change, it may be that other anti-immigrant actions by the Trump administration were responsible for some of the findings reported here.

Public Health Implications

A key implication of our findings is that rhetoric and the perceived threat of policy change are enough to produce chilling effects, prompting serious concern at further recent efforts targeting immigrants, such as eliminating sanctuary cities, family separation, and rescinding the Deferred Action for Childhood Arrivals program. Although most of these policies (including the public charge rule change) were challenged in court and were either not implemented or modified, it may be difficult to definitively determine their impact on immigrant well-being.

In the meantime, immigrant households, especially those with children, continue to experience higher levels of food insecurity.[9,32] Immigrants account for more than a quarter of the US population, and the health of the nation is inextricably linked to their well-being.[33] Absent efforts to systematically counteract the negative effects of rhetoric or policies that protect or restore access to public benefits, the utility of many national public health campaigns will likely be limited. AJPH

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CONTRIBUTORS

D. P. Miller and R. S. John conceptualized the study design and shared primary writing responsibilities. R. S. John led and D. P. Miller, M. Yao, and M. Morris contributed to data preparation and analysis. M. Yao and M. Morris contributed writing. All authors reviewed and approved the final version.

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CONFLICTS OF INTEREST

The authors do not have conflicts of interest from funding or affiliation-related activities.

HUMAN PARTICIPANT PROTECTION

This study used de-identified secondary data, so it was exempt from institutional review board review.

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Changes in the Public Charge Rule and Health of Mothers and Infants Enrolled in New York State's Medicaid Program, 2014–2019

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See also Alberto and Sommers, p. 1732.

Objectives. To examine the effect of the January 2017 leak of the federal government's intent to broaden the public charge rule (making participation in some public programs a barrier to citizenship) on immigrant mothers and newborns in New York State.

Methods. We used New York State Medicaid data (2014–2019) to measure the effects of the rule leak (January 2017) on Medicaid enrollment, health care utilization, and severe maternal morbidity among women who joined Medicaid during their pregnancies and on the birth weight of their newborns. We repeated our analyses using simulated measures of citizenship status.

Results. We observed an immediate statewide delay in prenatal Medicaid enrollment by immigrant mothers (odds ratio = 1.49). Using predicted citizenship, we observed significantly larger declines in birth weight (−56 grams) among infants of immigrant mothers.

Conclusions. Leak of the public charge rule was associated with a significant delay in prenatal Medicaid enrollment among immigrant women and a significant decrease in birth weight among their newborns. Local public health officials should consider expanding health access and outreach programs to immigrant communities during times of pervasive antiimmigrant sentiment. (*Am J Public Health*. 2022; 112(12):1747–1756. https://doi.org/10.2105/AJPH.2022.307066)

AJPH December 2022, Vol 112, No. 12

Since 1882, US immigration law has denied admission to people who are or are likely to become a public charge. The term public charge, however, was undefined until 1999, when regulatory guidance limited the definition to those who were primarily dependent on specific federal benefit programs for their income or requiring long-term institutionalized care.[1,2]

In 2017, the Trump administration indicated its intent to change the definition of public charge in a way that would constrain low-income immigrants' use of core public benefit programs essential to health and well-being. In January 2017, a draft executive order from the federal government to broaden the existing rule was leaked and circulated widely. A proposed rule was published in October 2018.[3] A final rule was issued in August 2019,[4] but its implementation was the subject of several court challenges. The rule ultimately went into effect briefly on February 24, 2020, though full implementation was stayed by the courts and after January 20, 2021, by the Biden administration.[5] On September 8, 2022, the Biden administration published a new set of rules that codifies the more generous pre–Trump era public charge guidance.[6]

When deemed a public charge, an individual is not eligible for lawful permanent resident (LPR) status, commonly known as holding a "green card," and will be denied entry or reentry to the United States. The rule does not directly affect other immigrants—those who already have LPR status, are naturalized US citizens, or are the citizen children of immigrants. In this article, we use the term "noncitizen" to refer to those without LPR status and the term "immigrant" to include all foreign-born persons.

The pre-2020 definition deemed immigrants a public charge when the use of cash assistance programs or government-funded institutionalized long-term care represented their primary source of economic support.[2] The new rule would have expanded this list of benefits by incorporating several public benefit programs that are widely used by low-income families and individuals to help meet basic needs, such as the Supplemental Nutrition Assistance Program (SNAP), Medicaid, and housing assistance, and would regard any use of these benefits as grounds for deeming an individual a public charge. In addition, the revised rule creates stricter income and wealth tests. The effects could have been substantial, because the use of these additional benefits is so widespread. While over the period 1997 to 2017 fewer than 3% of US-born citizens participated in the programs that comprised the criteria under the long-standing definition, nearly half (43% to 52%) participated in at least one of the programs that would have made them subject to the new public charge criteria had they been immigrants.[7,8] The proposed rule changes could have had far-reaching and direct effects on the composition, health, and economic stability of the targeted immigrant families. Because of confusion and fear of deportation or loss of future LPR status, "the chilling effect," they could have affected eligible immigrants who were not directly targeted by the rule but might nevertheless not enroll or renew public benefits for themselves or their (citizen) children. In addition, immigrants might not seek or might withdraw from public benefits that were not targeted by the rule, such as the Special Supplemental Nutrition Program for Women, Infants, and Children or Medicaid among pregnant women and children aged younger than 21 years.

Large-scale chilling effects caused by the new widened definition were reported broadly.[9–14] The Urban Institute found that 14.8% of adults in low-income families with children reported avoiding Medicaid or the Children's Health Insurance Program in 2019.[15] Research has also shown that potentially 2.1 million essential workers during the COVID-19 pandemic failed to enroll in Medicaid, and 1.3 million gave up SNAP because of concerns about the public charge rule.[9]

IMPORTANCE OF ACCESS TO HEALTH CARE

The impact of the public charge rule may be particularly consequential for the health of low-income pregnant immigrant women, who might delay Medicaid enrollment during pregnancy, which, in turn, could delay and reduce prenatal care utilization.[12,16] Lack of proper prenatal care during pregnancy might lead to lower birth weight and an increased likelihood of preterm birth. Health conditions such as maternal depression that go undiagnosed and untreated have been found to also negatively affect children's health, food security, and developmental outcomes.[17,18] Parental insurance coverage is associated with a greater likelihood that insured children have a usual source of health care and receive preventive services.[19–21] Studies have also shown that sociopolitical stressors, such as immigration raids and President Trump's inauguration, themselves significantly increased rates of preterm births and low birth weight.[22–24]

In the United States, 1 in 4 children live with at least 1 immigrant parent.[25] More than 10 million people live in immigrant families that receive 1 of the major public benefits that under newly proposed rules could be considered a "public charge."[8] This includes millions of US-born children with noncitizen parents. New York State (NYS) has one of the nation's largest immigrant populations; at 4.4 million people they constitute more than 20% of the state's total population. The contrast between New York City (NYC) and the suburban or rural areas in New York State also provides a unique opportunity to examine the effect of the public charge rule leak in urban versus nonurban areas.

Considering the importance of access to timely prenatal care for low-income immigrant women, the gaps in the literature, and the hostile environment that may be generated by antiimmigration policies and rhetoric, this study aimed to measure changes in Medicaid enrollment of pregnant low-income immigrant women as a result of the 2017–2020 public charge revisions.

METHODS

We used NYS Medicaid claims data for this analysis. The NYS Medicaid claims data include both fee-for-service claims and comprehensive managed care claims, which are of comparable quality.[26] The database includes Medicaid recipients' enrollment status, such as address history, demographic characteristics, and citizenship status, though the citizenship status variable is not available for those who joined Medicaid via the Health and Benefits Exchange Program after 2014. The database also includes detailed information on

Medicaid utilization, including date of service, diagnoses, and procedures.

Sample

We selected all infants born in NYS hospitals between September 2014 and December 2019. We then linked the infants to their mothers by using the Medicaid case number, infant's date of birth, and mother's hospital discharge date. On average, we identified more than 120 000 infants per year, and 89% were linked to their mothers (Appendix A, available as a supplement to the online version of this article at https://ajph.org). Our main sample was mothers who joined Medicaid during pregnancy (40%–48% of women who were pregnant each year) because NYS offers Medicaid to pregnant women at a relatively higher income threshold of $28 723 for a family of 1 regardless of immigration status.

Timing of the Public Charge Rule Impacts

We used January 2017 as the cutoff for the post period because the memo leaked during that time. We excluded pregnancies that had dates of birth between April and December 2016. Although January 2017 was the month of the inauguration and the leak of the memo, Trump announced his candidacy in June 2015 and gained popularity and large-scale media coverage from 2015 to 2016, so chilling effects may have already been triggered in this population before January 2017. We observed some evidence of the pre-2017 chilling effect in our data (Figures 1 and 2). We provided a set of sensitivity analyses including April to December 2016 in Appendix B (available as a supplement to the online version of this article at https://ajph.org).

Citizenship

We established 2 citizenship measures. For most, but not all, Medicaid beneficiaries, citizenship status is recorded in the enrollment record. The percentage of those without recorded immigration status increased by year, with 2019 the highest at 30%.

We cannot rule out that the missingness is not at random to the exposure of the public charge rule. To address those not reporting statuses, we used a conditional probability method to estimate a continuous measure that represents an individual's probability of being foreign-born conditioning on that individual's age (aged 18 years or older vs younger than 18 years), sex (male vs female), race (White, Black, Asian, Hispanic, and other), and census tract in NYS. Studies have used the American Community Survey (ACS) to examine the effect of the public charge rule by citizenship status.[11,13,14] We used data from the 2018 ACS Five-Year Estimate to construct a "sex by age by nativity and citizenship status" rate within each race/ethnicity group.

For instance, Person A is 20 years old, female, Hispanic, and living in a given census tract. To predict Person A's probability of being "foreign born," we use the estimate of the number of foreign-born people, and 18 years and older, female, and Hispanic living in that census tract as the numerator and the estimate of the total number of people who are 18 years and older, female, and Hispanic living in the given census tract as the denominator.

We verified the estimate using reported citizenship status. The estimate has a stronger predictive value outside NYC. For all reported noncitizens, the average predicted probability of being an immigrant in our model was 0.28 in NYC and 0.24 in the rest of the state; for all reported citizens, the average predicted probability of being an immigrant in our model was 0.15 in NYC and 0.04 in the rest of the state.

We included the predicted probability as a continuous variable ranging from 0 to 1 in all the regression models. In the time-series graphs only, we used a binary variable with 1 indicating a predicted probability between the third quartile and the maximum based on the distribution of known noncitizens (≥ 0.4 for NYC and ≥ 0.39 for the rest of the state) and 0 indicating a predicted probability between the minimum and the first quartile (≤ 0.13 for NYC and ≤ 0.02 for the rest of the state). We included a set of time-series graphs using the median (0.25 for NYC and 0.15 for the rest of the state) as the cutoff in Appendix C (available as a supplement to the online version of this article at https://ajph.org).

Outcomes

We evaluated delayed enrollment during pregnancy, prenatal care visits, low birth weight, and severe maternal morbidity (SMM). We used 2 measures of delayed Medicaid enrollment during pregnancy, after the end of the first trimester (≤ 6 months before birth), and after the end of the second trimester (≤ 3 months before birth). We evaluated whether mothers had any prenatal visits. Among those with at least 1 outpatient visit, we evaluated the change in the number of total visits and the days to the first outpatient visit since the imputed pregnancy date (280 days before the infant's date of birth). Low birth weight is a binary variable with 1 indicating 2500 grams or less. We

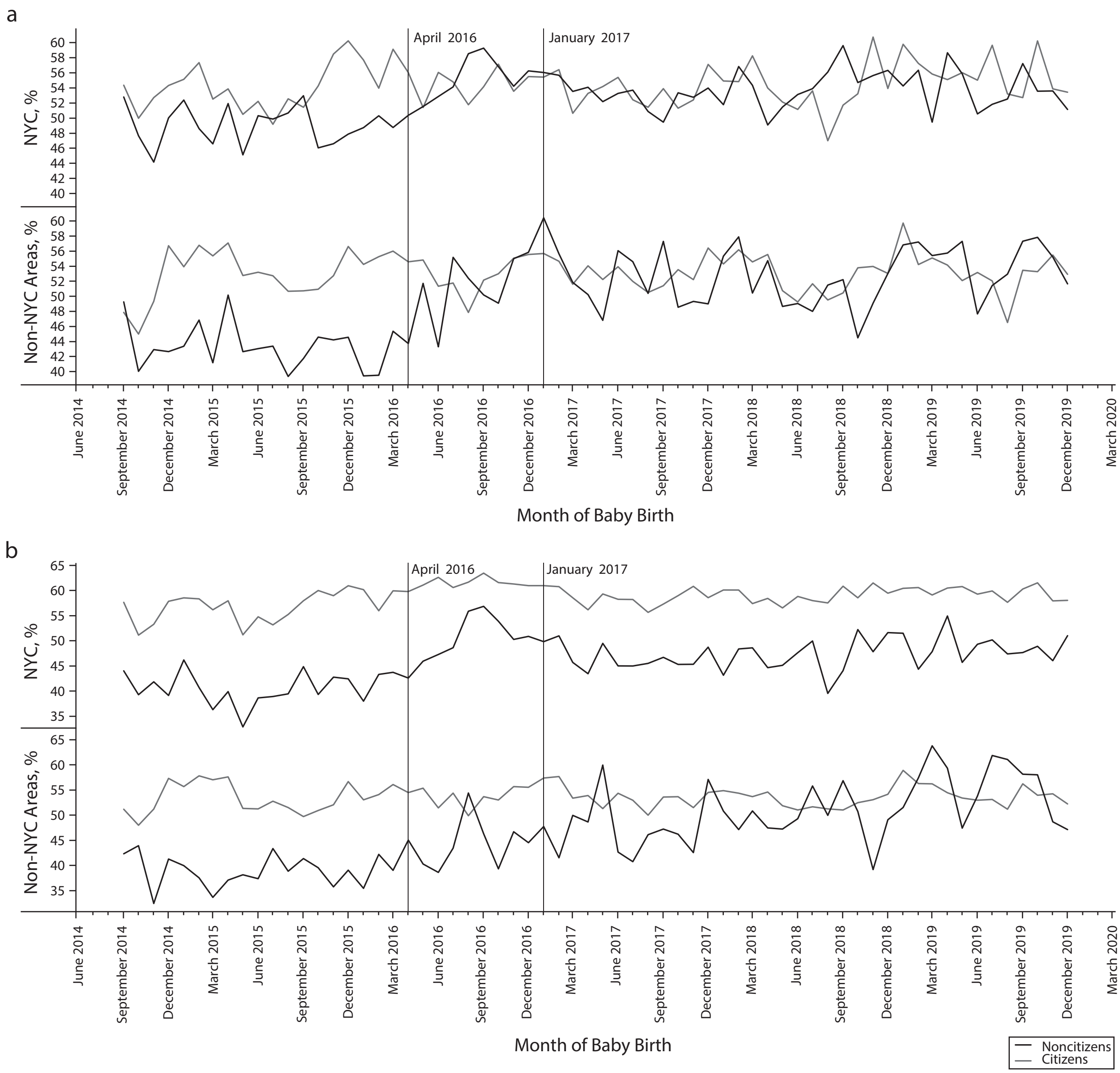

FIGURE 1— Percentage of Delayed Enrollment by (a) Reported Citizen Status and (b) Predicted Citizenship Status: New York State, 2014–2019

used the SMM definition provided by the Centers for Disease Control and Prevention. We included qualifying diagnoses or procedures for SMM-related inpatient visits 1 year after birth.

Statistical Analysis

We used SAS version 9.4 (SAS Institute, Cary, NC) to perform all statistical analyses. We used a comparative interrupted time-series (ITS) model and a difference-in-difference (DID) design to test for the immediate effect of the public charge rule. We then adjusted for the mother's age, race, county, and infant's birth month to control for individual, geographical, and seasonal effects (adjusted ITS). Lastly, we evaluated the overall effect of the public charge rule using a traditional DID model, post versus pre and noncitizens versus citizens, adjusting for age, race, county, and infant's birth month. We used logistic regression for all binary outcomes and linear regression for the continuous outcomes. We included the model statements in Appendix D (available as a supplement to the online version of this article at https://ajph.org).

In both sets of models, we used individual-level data. In reported citizenship models, we used citizen women

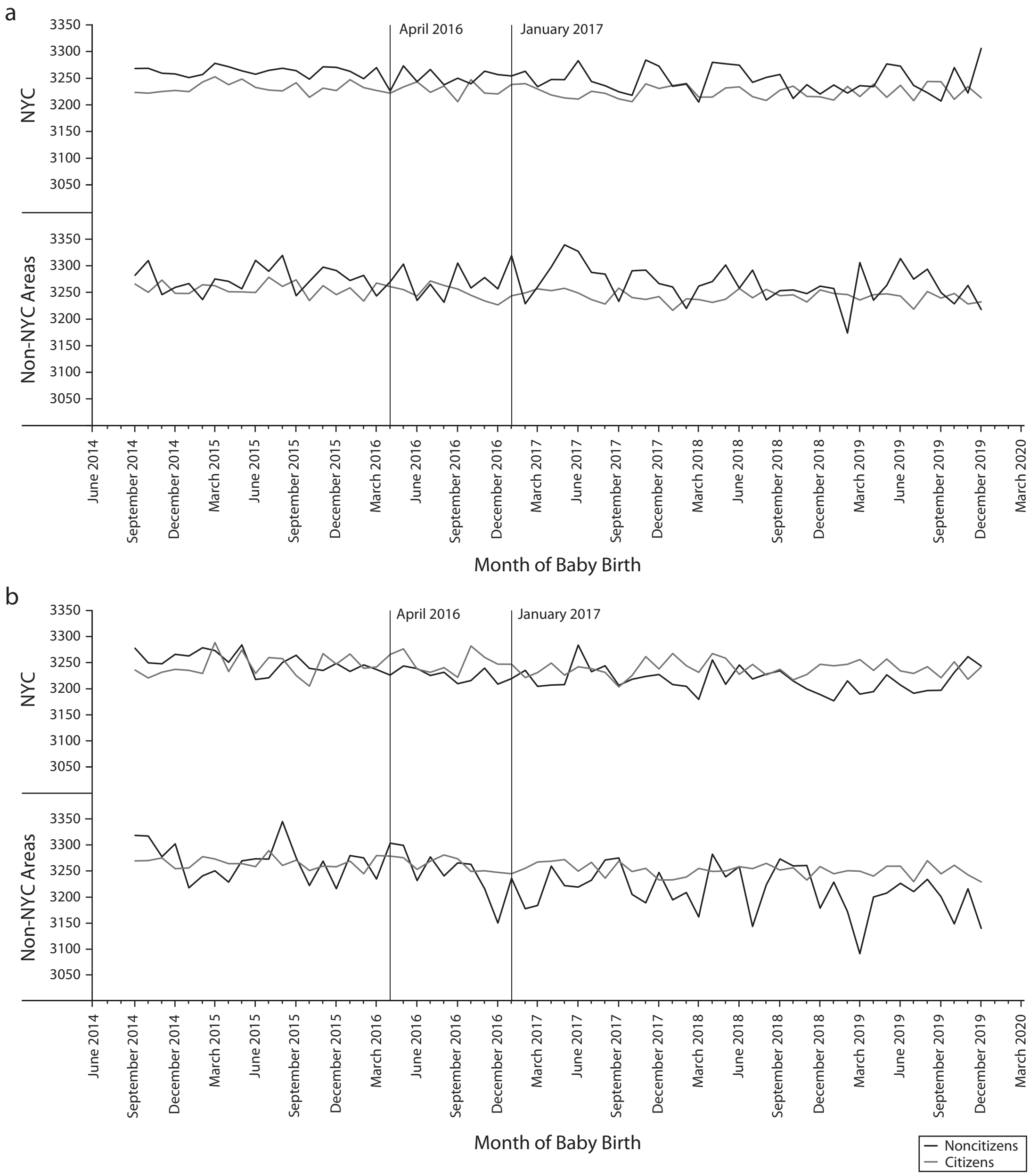

FIGURE 2— **Birth Weight by (a) Reported Citizen Status and (b) Predicted Citizenship Status: New York State, 2014–2019**

as the reference group; we excluded those in the unknown citizenship group. In the predicted citizenship models, we included individuals with both unknown and known citizenship; we used predicted probability for all individuals.

Sensitivity Analyses

We used 2 additional samples. The second sample included only the oldest child of the family (42%–44% each year) to account for the increased familiarity and comfort level with the Medicaid program (or belief that public charge status was already a given) at subsequent births. The third sample combined mothers who joined Medicaid before pregnancy with those who joined Medicaid during pregnancy (Appendix E, available as a supplement to the online version of this article at https://ajph.org).

We also looked at the immediate and overall effects of all the outcomes among Hispanics, Asians, and unknown racial groups (Appendix F, available as a supplement to the online version of this article at https://ajph.org).

RESULTS

Table 1 shows the demographic and outcome distributions by reported

TABLE 1— Means and Percentages of Demographic and Outcome Variables by Reported Citizenship and Geography Among Pregnant Women Enrolled in the New York State Medicaid Program, 2014–2019

	NYC US Citizens		NYC US Noncitizens		Non-NYC US Citizens		Non-NYC US Noncitizens	
	Pre	Post	Pre	Post	Pre	Post	Pre	Post
No.	15 096	15 527	23 362	17 037	25 347	33 158	7 256	6 921
Age, y (mean)	27	28	30	30	27	28	30	31
Hispanic, %	14.19	9.46	37.66	22.92	6.30	3.49	41.45	12.94
Non-Hispanic, %								
Asian	6.12	5.97	14.81	18.31	2.51	2.60	10.17	15.51
Black	28.90	29.05	18.61	15.44	18.00	18.69	10.58	13.46
White	15.44	15.54	9.36	9.77	48.53	47.46	8.53	10.70
Other	6.80	6.22	5.55	6.08	4.00	4.06	5.11	7.00
Unknown race	28.51	33.72	13.99	27.46	20.63	23.67	24.14	40.35
Outcome measures								
Medicaid enrollment delays, %								
≤ 6 mo	54	55	49	54	53	53	43	53
≤ 3 mo	21	22	22	20	21	21	12	18
Prenatal visits, any, %	90	91	95	96	91	92	97	96
Among those with any, no. of visits, mean	8	8	9	9	8	8	11	10
Among those with any, no. of days to first visit, mean	130	132	133	137	127	131	115	132
Birth weight, grams	3 194	3 187	3 259	3 243	3 252	3 235	3 273	3 239
Low birth weight, %	9	9	7	7	8	9	6	7
Severe maternal morbidity, %	5	7	6	6	4	4	4	5

Note. NYC = New York City.

citizenship status. Reported noncitizens and citizens were similar in age. Noncitizens were more likely to report Hispanic or Asian race/ethnicity.

Delayed Enrollment

We observed both immediate and overall effects of the public charge rule on delayed Medicaid enrollment (Table 2). The adjusted ITS model results showed increased delayed enrollment immediately after January 1, 2017, in NYS using both measures of citizenship. In NYS, delayed enrollment (≤ 6 months) increased (odds ratio [OR] = 1.49; 95% confidence interval [CI] = 1.26, 1.77) comparing noncitizens to citizens in the immediate post–public charge period.

The overall effect (DID) in delayed enrollment (≤ 6 months) in NYS had an OR of 1.16 (95% CI = 1.09, 1.23). As the predicted probability of being an immigrant increased from 0 to 1, immediate delayed enrollment (≤ 6 months) increased (OR = 1.89; 95% CI = 1.33, 2.69), while overall the odds of delayed enrollment increased (OR = 1.43; 95% CI = 1.27, 1.61). The large increase is driven by the upstate New York and Long Island (non-NYC) area. In NYC, delayed enrollment (≤ 6 months) comparing noncitizens to citizens in the post–public charge period was OR = 1.36 (95% CI = 1.09, 1.70) for the immediate delay and OR = 1.07 (95% CI = 0.99, 1.16) for the overall delay, while in non-NYC areas, it was OR = 1.94 (95% CI = 1.34, 2.82) for the immediate delay and OR = 1.53 (95% CI = 1.34, 1.75) for the overall delay.

As the predicted probability of being an immigrant increased from 0 to 1, the OR of immediate delayed enrollment (≤ 6 months) in NYC was positive, but not statistically significant (OR = 1.44; 95% CI = 0.88, 2.36), while overall delayed enrollment increased (OR = 1.29; 95% CI = 1.09, 1.52). We observed significant immediate and overall delays using predicted citizenship (≤ 6 months) in non-NYC areas: immediate OR = 2.54 (95% CI = 1.43, 4.51); overall OR = 1.37 (95% CI = 1.11, 1.68).

We observed a significant overall increase in extremely delayed Medicaid enrollment (≤ 3 months) during

TABLE 2— The Immediate and Overall Effects of the Public Charge Rule on Maternal and Child Health Outcomes, Comparing Noncitizens and Predicted Immigrants to Citizens: New York State, 2014–2019

	Reported Citizenship			Predicted Citizenship		
	NYS, Estimate (95% CI)	NYC, Estimate (95% CI)	Non-NYC, Estimate (95% CI)	NYS, Estimate (95% CI)	NYC, Estimate (95% CI)	Non-NYC, Estimate (95% CI)
Immediate effect: comparative interrupted time series						
Medicaid enrollment delays						
≤6 mo, OR	1.49 (1.26, 1.77)	1.36 (1.09, 1.70)	1.94 (1.34, 2.82)	1.89 (1.33, 2.69)	1.44 (0.88, 2.36)	2.54 (1.43, 4.51)
≤3 mo, OR	1.16 (0.94, 1.42)	1.03 (0.79, 1.35)	1.51 (0.89, 2.55)	0.96 (0.61, 1.51)	0.83 (0.45, 1.52)	2.18 (0.96, 4.95)
Prenatal visits, OR	0.70 (0.49, 1.00)	0.59 (0.38, 0.90)	0.58 (0.21, 1.57)	0.91 (0.40, 2.05)	0.43 (0.14, 1.32)	1.16 (0.27, 5.05)
No. of visits, b	−0.08 (−0.60, 0.43)	0.13 (−0.57, 0.82)	−1.47 (−2.55, −0.39)	−0.07 (−1.15, 1.00)	0.41 (−1.12, 1.94)	−2.13 (−3.80, −0.46)
Days to first visit, b	15.65 (9.80, 21.49)	11.18 (3.58, 18.79)	25.57 (12.67, 38.46)	26.33 (14.24, 38.43)	14.19 (−2.48, 30.87)	48.53 (28.73, 68.34)
Birth weight, grams, b	2.93 (−41.61, 47.48)	25.51 (−32.18, 83.19)	−19.33 (−118.69, 80.03)	−22.09 (−112.80, 68.62)	−57.00 (−181.17, 67.17)	8.55 (−142.27, 159.38)
Low birth weight, OR	1.25 (0.91, 1.73)	1.18 (0.79, 1.77)	1.48 (0.69, 3.16)	0.99 (0.49, 2.00)	0.99 (0.38, 2.59)	1.03 (0.33, 3.19)
SMM, OR	1.21 (0.83, 1.77)	1.49 (0.95, 2.34)	1.13 (0.41, 3.07)	1.10 (0.48, 2.50)	1.92 (0.68, 5.45)	0.41 (0.10, 1.77)
Overall effect: difference-in-difference[a]						
Medicaid enrollment delays						
≤6 mo, OR	1.16 (1.09, 1.23)	1.07 (0.99, 1.16)	1.53 (1.34, 1.75)	1.43 (1.27, 1.61)	1.29 (1.09, 1.52)	1.37 (1.11, 1.68)
≤3 mo, OR	0.89 (0.83, 0.96)	0.78 (0.71, 0.85)	1.63 (1.35, 1.96)	1.28 (1.09, 1.49)	1.14 (0.93, 1.40)	1.87 (1.41, 2.49)
Prenatal visits, OR	0.85 (0.75, 0.96)	0.86 (0.74, 1.00)	0.58 (0.41, 0.81)	0.70 (0.53, 0.93)	0.77 (0.53, 1.13)	0.79 (0.47, 1.32)
No. of visits, b	−0.22 (−0.39, −0.04)	−0.13 (−0.36, 0.11)	−0.57 (−0.97, −0.18)	−0.75 (−1.12, −0.39)	−0.37 (−0.88, 0.14)	−0.94 (−1.55, −0.34)
Days to first visit, b	1.86 (−0.16, 3.87)	−0.70 (−3.30, 1.90)	18.96 (14.24, 23.68)	13.84 (9.76, 17.92)	7.16 (1.62, 12.70)	26.71 (19.55, 33.88)
Birth weight, grams, b	−6.83 (−22.22, 8.56)	−0.17 (−19.94, 19.60)	−37.08 (−73.31, −0.86)	−55.96 (−86.57, −25.35)	−52.10 (−93.45, −10.75)	−91.42 (−145.82, −37.01)
Low birth weight, OR	1.04 (0.93, 1.17)	1.10 (0.96, 1.26)	0.99 (0.76, 1.30)	1.09 (0.86, 1.38)	1.15 (0.84, 1.58)	1.20 (0.80, 1.81)
SMM, OR	1.04 (0.91, 1.18)	0.90 (0.77, 1.05)	1.65 (1.15, 2.36)	0.60 (0.45, 0.80)	0.34 (0.24, 0.49)	1.31 (0.77, 2.23)

Note. CI = confidence interval; NYC = New York City; NYS = New York State; OR = odds ratio; SMM = severe maternal morbidity.

[a]We evaluated the overall effect of the public charge rule using a traditional difference-in-difference model—post vs pre and noncitizens vs citizens, adjusting for age, race, county, and infant's birth month.

pregnancy outside NYC (OR = 1.63; 95% CI = 1.35, 1.96) using reported citizenship and OR = 1.87 (95% CI = 1.41, 2.49) using predicted citizenship.

Prenatal Care Visits

The results showed a significant and overall decrease in the fraction of mothers who had prenatal visits in NYS (OR = 0.85; 95% CI = 0.75, 0.96) using reported citizenship and OR = 0.70 (95% CI = 0.53, 0.93) using predicted citizenship.

Among those with prenatal care visits, we observed decreases in the number of visits and delays to the first visit both immediately and overall. The effect was driven by non-NYC areas: using reported citizenship, mothers immediately had 1.47 (95% CI = −2.55, −0.39) fewer prenatal visits and delayed 25.57 days (95% CI = 12.67, 38.46), and 0.57 (95% CI = −0.97, −0.18) fewer visits and 18.96 (95% CI = 14.24, 23.68) days in the delay overall. Using predicted citizenship, compared with nonimmigrant mothers in non-NYC areas, immigrant mothers had 2.13 (95% CI = −3.80, −0.46) fewer prenatal visits and experienced 48.53 (95% CI = 28.73, 68.34) days in the delay to the first prenatal visit immediately and 0.94 (95% CI = −1.55, −0.34) visits and 26.71 (95% CI = 19.55, 33.88) days overall.

Low Birth Weight

We observed significant overall decreases in birth weight in non-NYC areas: newborns of reported noncitizen mothers weighed 37.08 grams less (95% CI = –73.31 grams, −0.86 grams) than those of citizen mothers; newborns of predicted immigrant mothers weighted 91.42 grams less (95% CI = −145.82 grams, −37.01 grams). We did not observe significant changes in the prevalence of low birth weight using the cutoff of 2500 grams or less in the main analyses.

Severe Maternal Morbidity

Compared with reported citizens, the overall odds of SMM for noncitizens increased (OR = 1.65; 95% CI = 1.15, 2.36) in the post period. Using predicted citizenship, we observed significant decreases in SMM in NYC (OR = 0.34; 95% CI = 0.24, 0.49), as well as in NYS as a whole (OR = 0.6; 95% CI = 0.45, 0.8). We did not observe significant immediate effects in SMM using either reported or predicted citizenship.

Sensitivity Analyses

Both the oldest child sample and the all-mothers sample showed significant and immediate delayed enrollment (≤ 6 months) and delays to the first prenatal visit (Appendix E). We observed significant and consistent overall effects of delayed enrollment (≤ 6 months) in the all-mothers sample.

We observed significant effects for both immediate and overall delayed enrollment (≤ 6 months) among Asians using predicted citizenship. We observed positive but not statistically significant results for immediate delayed enrollment among Hispanics using both measures of citizenship (Appendix F). Among those of unknown race, we observed significant statewide overall effects for both measures of delayed enrollment, the number of prenatal visits, days to the first visit, and SMM for outside NYC only (Appendix F).

DISCUSSION

We found that the public charge rule was associated with large and significant damage to the health of immigrant mothers and children in the month of the memo leak, 3 years before it went into effect. In a way, the early timing of our study is evidence of a broader chilling effect beyond the public charge rule—the longer-standing generalized fear among immigrants about seeking public supports given pervasive antiimmigrant sentiment and racial biases that were stoked by the Trump administration.

Among studies and reports that directly examined the effect of the public charge rule on health care, various timing and data sources have been used to define the post–public charge period. The set of reports from the Urban Institute looked at Internet surveys conducted in December 2018 to 2020.[15,27–29] One study used ACS survey data that compared annual Medicaid and SNAP enrollment changes from 2016 to 2019.[14] Other studies based on population surveys and provider surveys looked at effects in 2019.[11,12] We found that 1 study examined changes from August 2016 to June 2019 using SNAP administrative program data, although the DID effect was estimated as of September 2018.[30] Compared with these studies, our study used individual-level administrative data on Medicaid program use and estimated the direct and significant effect at the earliest timing, in January 2017.

We observed a statewide effect in delayed Medicaid enrollment. The magnitude of such delay is substantial. Among all noncitizen mothers who joined Medicaid during pregnancy, 48% joined in the second trimester or later in March 2016, compared with 57% in January 2017. Similarly, among mothers who lived in areas with higher percentages of noncitizens, 42% joined Medicaid in the 2nd trimester or later in March 2016 versus 49% in January

2017. While Medicaid receipt by pregnant immigrant women would not, under the rule, be considered in a public charge determination,[31] declines in Medicaid coverage could occur beyond those directly targeted by the rule.

Our outcomes for prenatal care are consistent with reports indicating that immigrant women were afraid to get prenatal care because of fear of the public charge rule.[32] The Kaiser Family Foundation found that half of the health centers surveyed reported a decline in health care use by immigrant patients, especially immigrant pregnant women who were not enrolling in or were disenrolling from Medicaid out of fear of the consequences of being deemed a public charge.[12]

The literature has shown that immigrants can have different experiences of the system within the same state.[33] We have seen evidence of this variability in our study. One such variation between NYC and the rest of the state is that NYC has done extensive outreach to the immigrant communities about seeking care and health services, partnering with dozens of community-based organizations and the public hospital system,[34] in addition to laws that NYS as a whole has put in place to support immigrants including those who are undocumented.[35,36]

For all outcome measures, we observed worse outcomes outside NYC areas. This may be, in part, because we were better able to predict citizenship outside NYC. Even using measured citizenship, however, we observed a larger (and statistically significant) reduction in the number of prenatal visits outside NYC. Among those with reported citizenship, the estimated delay in seeking prenatal care was about 18 days for noncitizen mothers, contributing to a significant reduction in the total number of prenatal visits. Together with the significant delay and reduction in prenatal care, the odds of SMM increased significantly; birth weight also decreased significantly, by about 37 grams in the post period. We did not observe significant effects of any of the mentioned results in NYC.

Strengths and Limitations

Some of the strengths of the study included the use of large-scale claims data at the individual level that allowed us to study the universe of low-income pregnant women on Medicaid and measure nuanced enrollment and health outcomes for both individual infants and mothers.

As a limitation, we had many unreported citizenships in the data, which could threaten the validity of the study by introducing selection bias. We addressed the limitation by estimating the effects using predicted citizenship. Because we only looked at NYS, generalizing the results to states with different Medicaid or immigration policies would be another limitation.

Public Health Implications

Our study demonstrated that the rule changes the Trump administration proposed had far-reaching chilling effects on the health of immigrant mothers and their (citizen) infants. We found larger effects in suburban and rural areas, perhaps because advocacy and community resources are less available in such areas. Local public health officials should consider expanding health access and outreach programs to immigrant communities during times of pervasive antiimmigrant sentiment. AJPH

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CONFLICTS OF INTEREST

Authors have no conflicts of interest to disclose.

HUMAN PARTICIPANT PROTECTION

The institutional review board at New York University has exempted the study (IRB-FY2018-1285).

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Use of Law by US States During the COVID-19 Pandemic With Respect to People Who Were Undocumented

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See also Bustamante et al., p. 1729.

Objectives. To systematically identify and analyze US state-level legislation concerning people who were undocumented during the COVID-19 pandemic, from January 2020 through August 2021.

Methods. Using standard public health law research methods, we searched Westlaw's online database between November 2021 and January 2022 to identify legislation addressing COVID-19 and people who were undocumented. We abstracted relevant information, analyzed the data, and identified primary themes for each bill and resolution.

Results. Sixty-six bills and resolutions, from 13 states, met the inclusion criteria. Legislation addressed 5 primary themes: eligibility and access to health-related services (n = 16), health and personal information (n = 10), housing assistance (n = 13), job security and employment benefits (n = 14), and monetary assistance (n = 13).

Conclusions. Approximately one quarter of state legislatures introduced bills or resolutions regarding people who were undocumented and COVID-19. State-level laws are an important tool to mitigate the disproportionate impact of public health emergencies on vulnerable groups.

Public Health Implications. As states shift attention away from the exigencies of COVID-19, this research provides insight into how law might be used to protect those who are undocumented throughout the full cycle of future public health emergencies. (*Am J Public Health*. 2022;112(12): 1757–1764. https://doi.org/10.2105/AJPH.2022.307090)

While the federal government has passed high-profile legislation to control the spread of COVID-19 and mitigate its economic impact—including the American Rescue Plan Act[1]; Coronavirus Aid, Relief, and Economic Security (CARES) Act[2]; and Families First Coronavirus Response Act[3]—US state governments have also played a major role in pandemic response. For example, governors have issued and refined states of emergency, stay-at-home orders, mask mandates, and quarantine guidance.[4] States also took significant administrative action, such as expanding conditions that qualify for emergency Medicaid (i.e., coverage for treatment of emergency medical conditions).[5] Legislatures in all 50 states and the District of Columbia have introduced a host of COVID-19–related bills tailored to their populations' needs.[6] This frenzy of legislative activity provides insight about policymaker priorities and reveals the range of ways that states might use law to protect their most vulnerable populations during an infectious disease emergency.

The United States has experienced a higher COVID-19 death rate compared with other well-resourced and similarly sized countries,[7] with a disproportionate share of deaths experienced by certain populations. For example, Black and Hispanic people in the United States were 2 times more likely than White people to die from COVID-19.[8] The pandemic has had a disparate impact on people who were undocumented (i.e., immigrants residing in the United States without official government authorization), with these communities experiencing disproportionately high rates of COVID-19 morbidity and mortality.[9] People who were

AJPH December 2022, Vol 112, No. 12

undocumented were especially vulnerable to COVID-19, in part because of barriers accessing health care and exclusion from federal stimulus payments issued during the pandemic.[10,11] More than 70% of the approximately 7 million undocumented workers in the United States are direct service workers who cannot work remotely, placing them at higher risk of COVID-19 exposure and infection.[10,12] In addition, the COVID-19 pandemic has been associated with significant anti-immigrant sentiment, often directed at people who are undocumented.[13,14] For example, some Republican lawmakers falsely tried to attribute COVID-19 surges in the South to migrants who crossed the southern border.[15] Law can be an important tool to protect this vulnerable population, especially during a public health emergency like the COVID-19 pandemic.

This article presents a legal mapping study[16] that systematically identified and analyzed US state-level legislative activity related to people who were undocumented during the COVID-19 pandemic, from January 1, 2020, through August 31, 2021. Proposed and enacted bills and resolutions were included, with accompanying analysis of their objectives and implications for future public health policy.

METHODS

Using standard public health law research methods,[16] we identified proposed and enacted state-level legislation related to COVID-19 and people who were undocumented. The search was conducted using the Westlaw database between November 2021 and January 2022. We used standardized search terms to identify bills and resolutions from all 50 US states and Washington, DC, that were introduced between January 1, 2020, and August 31, 2021. The start date was chosen because the first confirmed COVID-19 case in the United States occurred in Washington State in January 2020.[17] The end date was chosen because it allowed us to capture legislation proposed during the 18 months following March 2020—the month when the World Health Organization (WHO) declared a global pandemic and the United States declared a nationwide emergency.[18,19] For purposes of our analysis, we recorded the status of each bill or resolution through February 28, 2022, to account for a 6-month window of legislative activity following the introduction of bills and resolutions within our designated timeframe.

Our final search string comprised 2 sets of standardized search terms: (1) terms related to COVID-19 and (2) terms related to people who were undocumented. Initial search terms were generated using the research team's a priori knowledge and preliminary online research to understand the language used to describe our topics of interest. Next, we reviewed examples of relevant bills and used an iterative process to identify additional applicable search terms. University law librarians were consulted to help format each term (e.g., with connectors) to ensure that the search returned the maximum number of relevant bills and resolutions.

A public health law expert (J. R.) with relevant subject matter expertise reviewed the final search string. The final search terms, which used Boolean terms and connectors, were (COVID! OR coronavirus OR "corona virus" OR SARS-CoV-2 OR SARS OR pandemic OR outbreak! OR epidemic OR "health emergenc!" OR "infectious disease" OR quarantin! OR isolat! OR "social distanc!" OR "personal protective equipment" OR PPE OR mask! OR "face covering!" OR ventilat!) AND (immigrant! OR immigrat! OR undocumented OR migrant! OR migrat! OR alien! OR "foreign born!" OR foreign-born! OR "foreign national!" OR "unauthorized person!" OR noncitizen! OR nonresident! OR refuge! OR asyl! OR deport! OR mexic! OR spanish! OR hispanic! OR latin! OR visa! OR "green card!" OR "resident card!" OR DACA OR DAPA OR citizenship OR "national origin").

Some of the undocumented-specific terms in the search string are derogatory (e.g., foreign-born, alien). This terminology does not reflect the views or the lexicon used by the authors. We included search terms with this language to maximize the number of relevant bills and resolutions captured.

A research team member (E. D.) conducted a preliminary screen of each bill or resolution that our search query yielded. The initial scan involved reviewing the search terms within each piece of legislative text to understand whether the bill or resolution met inclusion criteria (i.e., pertained to people who are undocumented in the context of COVID-19). A second research team member (L. R.) then reviewed the full text of each bill or resolution identified in the initial screening process to determine whether they should remain in the final data set. When disagreement arose regarding inclusion or exclusion of certain proposed legislation, 2 team members reviewed the text together, discussed any points of disagreement, and reached a determination by consensus.

Where there were multiple versions of a bill or resolution, we removed duplicates and retained the most recent version in the final data set. For instances in which there were cross-listed versions of the same bill or resolution with the same date, we retained the legislation from the state's higher chamber in the final data set. We only included proposed and enacted bills and resolutions in the final data set, and excluded other types of documents (e.g., legislative memos). We also excluded legislation if language pertaining to the research question was found solely in prefatory sections, such as the preamble or legislative intent and findings. For each bill or resolution, we abstracted information on jurisdiction, bill or resolution number, date of introduction, bill or resolution sponsor(s) and political party, status of the bill or resolution, primary theme addressed, and whether the bill or resolution had the potential to be beneficial, harmful, or neutral toward people who are undocumented.

We summarized information whenever possible with descriptive statistics. We repeatedly reviewed each bill and resolution to determine the primary topic or theme it addressed. This included multiple rounds of review and comparison of bill or resolution text and discussion among research team members. For bills and resolutions that potentially addressed more than 1 theme, we reviewed the relevant language and determined a "primary theme" based on the topic that was most frequently or prominently addressed. Thus, we categorized each bill or resolution into only 1 theme.

RESULTS

The search yielded 5344 pieces of proposed and enacted legislation. Sixty-six bills and resolutions satisfied our inclusion criteria and were included in the final data set (Table A, available as a supplement to the online version of this article at https://ajph.org). Included bills and resolutions came from 13 states and were introduced between April 2020 and June 2021 (Table A and Table 1). Of the 13 states where legislation was introduced, 12 had Democratic-controlled legislatures and 1 had a split legislature (Table 1).

One resolution was adopted and 16 bills were passed (Table 1). Of the 66 bills and resolutions, 62 (94%) were potentially protective or beneficial toward people who were undocumented (e.g., expanding eligibility for stimulus payments) and 4 (6%) were neutral (e.g., creating a task force or allocating funds for an assessment of

TABLE 1— Bills or Resolutions Related to COVID-19 and People Who Were Undocumented by State of Introduction, Bill Status, and Select State Characteristics: United States, January 2020–August 2021

	Bill and Resolution Status			Select State Characteristics		
State	Total No. Introduced	Passed No. (%)	Did Not Pass[a] No. (%)	Undocumented % of State Population[20,b]	Legislature Partisan Control[21]	Political Party of Governor[21]
California	17	7 (41.2)	10 (58.8)	5.6	Democrat	Democrat
Hawaii	1	0 (0)	1 (100)	3.3	Democrat	Democrat
Illinois	8	4 (50.0)	4 (50.0)	3.2	Democrat	Democrat
Massachusetts	2	0 (0)	2 (100)	3.8	Democrat	Republican
Minnesota	4	0 (0)	4 (100)	1.7	Split	Democrat
Nevada	2	0 (0)	2 (100)	7.1	Democrat	Democrat
New Jersey	4	0 (0)	4 (100)	5.2	Democrat	Democrat
New York	13	2 (15.4)	11 (84.6)	3.6	Democrat	Democrat
Oregon	2	0 (0)	2 (100)	2.6	Democrat	Democrat
Rhode Island	2	0 (0)	2 (100)	2.8	Democrat	Democrat
Vermont	1	1 (100)	0 (0)	0.1	Democrat	Republican
Virginia	1	1 (100)	0 (0)	3.4	Democrat	Democrat
Washington	9	2 (22.2)	7 (77.8)	3.3	Democrat	Democrat
Total	66	17 (25.7)	49 (74.3)	. . .	. . .	. . .

[a]The "did not pass" number includes 5 bills that were introduced with language that addressed individuals who were undocumented within the context of the COVID-19 pandemic. However, this language was dropped from the bills as they worked their way through the legislative process. Versions of these bills—without language of relevance to the research question—were passed. Because of the removal of relevant language, these bills are categorized as "did not pass" for purposes of this study.
[b]Information based on 2016 estimates.

COVID-19 impact). Bills and resolutions were categorized into 5 themes:

1. eligibility and access to health-related services,
2. health and personal information,
3. housing assistance,
4. job security and employment benefits, and
5. monetary assistance (Table 2).

Eligibility and Access to Health-Related Services

Sixteen bills (24.2%) pertained to eligibility and access to health-related services. Of these, 4 (25%) became law. The most prevalent type of bill (7/16; 43.8%) within this theme proposed expansion of access to health care coverage and medical services during the pandemic for people who were undocumented. For example, the Illinois legislature enacted a bill that would temporarily expand coverage for treatment related to COVID-19 via the Illinois Department of Health for individuals who were not US citizens, including those who were undocumented (Illinois SB 2294 [2021]). Four bills would make undocumented individuals eligible for nonmedical COVID-19–specific services. For example, Virginia's governor approved a bill that would classify COVID-19–related testing, treatment, and vaccination as "emergency services" that are extended to certain individuals not lawfully admitted for permanent residence in the United States (Virginia HB 2124 [2021]). Two proposed bills from California had the goal of expanding food assistance to people regardless of immigration status (California AB 221 [2021], California SB 464 [2021]). The final 3 bills in this category sought to expand availability of mental health services (Oregon HB 2949 [2021]), appropriate funds for a study on service access (Washington SB 5091 [2021]), and create a task force (Washington HB 1340 [2021]).

Health and Personal Information

Nine bills and 1 resolution (15.2%) addressed health and personal information. Of these, 1 (10%) became law. Most of these bills (9/10; 90%) sought to prevent immigration authorities from accessing information related to contact tracing, vaccine status, testing, or other COVID-19 health data. The enacted bill, passed in New York, prevents immigration authority personnel from serving as contact tracers (New York SB 900 [2021]). Unsuccessful bills, such as 1 from Washington State, would broadly prohibit usage of COVID-19–related health data for purposes of immigration or law enforcement (Washington HB 1127 [2021]).

Housing Assistance

Thirteen bills (19.7%) sought to extend housing assistance during the COVID-19 pandemic. Each bill included the goal of increasing access to rental assistance for people who were undocumented. Of these, 4 (30.8%) were passed into law. For example, the California legislature passed a bill that allowed all persons—regardless of immigration status—to apply for rental assistance, and it prevents landlords from reporting or threatening to report a tenant to immigration authorities (California AB 832 [2021]). Four bills also included utility assistance as a housing-related benefit for which people who were undocumented may be eligible. For example, the New York legislature enacted a bill that specifies that households—regardless of immigration status—are eligible for rental assistance, utility assistance, or both (New York AB 3006 [2021]).

Job Security and Employment Benefits

Twelve bills and 2 resolutions (21.2%) pertained to job security and employment-related benefits. Of these, 2 bills and 1 resolution (21.4%) were passed. Four bills sought to prevent employers from taking "retaliatory personnel action" against employees, including reporting or threatening to report immigration status. Four bills or resolutions had the goal of providing legal documentation to workers. For example, legislators in Massachusetts introduced a bill that would expand eligibility for state licensure (e.g., driver's license, identification card) to people who did not qualify for a social security number, and the bill specified that people would not be asked about citizenship or immigration status during the application process (Massachusetts SB 2289 [2021]). The other 3 documentation-related bills and resolutions sought to expand undocumented workers' access to work permits (California AB 1510 [2021]), visas and green cards (New Jersey AR 196 [2020]), and residency status (Illinois SR 100 [2021]).

Two bills from California—1 proposed and 1 enacted—sought to provide grants to small businesses, including those owned and operated by undocumented individuals (California AB 151 [2021], California SB 151 [2021]). Two bills would provide benefits to people who became unemployed because of the COVID-19 pandemic. For example, the California legislature passed into law a bill that provided education and training grants for those who lost employment during the pandemic, and the law specifies that people who were

TABLE 2— Status of State-Level Bills and Resolutions by Theme and Examples Within Each Theme: United States, January 2020–August 2021

Introduced No. (%)	Passed No. (%)	Did Not Pass[a] No. (%)	Examples: Bill or Resolution Number (State): Description of Bill or Resolution
Eligibility and access to health-related services			
16 (24.2)	4 (25.0)	12 (75.0)	HB 2124 (Virginia): Declares that COVID-19 testing, treatment, and vaccines be considered “emergency services,” meaning these services are extended to certain individuals not lawfully admitted for permanent residence in the United States. SB 1620 (New York): Would offer free COVID-19 testing to all uninsured people, regardless of immigration status. SB 1515 (Massachusetts): Would appoint a “director of COVID-19 vaccination equity and outreach” to address barriers to vaccination that disproportionately affect marginalized communities including people who are undocumented.
Health and personal information			
10 (15.2)	1 (10.0)	9 (90.0)	HB 3120 (Illinois): Would ensure information collected by contact tracers remains confidential and would bar sharing this information with law enforcement or immigration authorities. Would prohibit law enforcement and immigration authorities from being contact tracers. HB 1127 (Washington): Would protect against usage of COVID-19 health data for law enforcement or immigration purposes. Would prohibit COVID-19 health data from being disclosed to or collected by law enforcement or immigration authorities. SB 6541 (New York): Would prohibit vaccine navigators, vaccine providers, and immunity passport providers from providing personal health information to law enforcement or immigration authorities.
Housing assistance			
13 (19.7)	4 (30.8)	9 (69.2)	SB 668 (Rhode Island): Would forgive rental and mortgage payments during a declared health emergency, including the COVID-19 pandemic. Would specify that “affordable housing operators” cannot refuse to rent based on an individual’s identity, including immigration status. SB 91 (California): Creates an emergency rental assistance program to provide funds for rent and utilities to individuals who have been financially impacted by the COVID-19 pandemic. Specifies that all people—regardless of citizenship or immigration status—can apply for the assistance. SB 668 (Illinois): Would create the “COVID-19 Federal Emergency Rental Assistance Program Act” to administer federal funds to help individuals cover the cost of rent payments. Would specify that unless otherwise necessary to comply with the law, program eligibility shall not consider applicants’ immigration status.
Job security and employment benefits			
14 (21.2)	3 (21.4)	11 (78.6)	SF 1518 (Minnesota): Would establish the Essential Workers Emergency Leave Act to provide paid sick leave to employees who were not eligible for funds under the federal Families First Coronavirus Response Act. Would prevent employers from taking “retaliatory personnel action” against employees who request or receive emergency paid sick leave by prohibiting actions such as disclosing or threatening to disclose an employee’s immigration status. SB 151 (California): Establishes the California Microbusiness COVID-19 Relief Program to provide grants to eligible small businesses that have been negatively impacted by the pandemic. The bill specifies that people who are undocumented are eligible to receive micrograntsand eligibility determinations will not require information about an individual’s immigration status. SB 5438 (Washington): Would provide unemployment benefits to workers who lost their jobs because of the COVID-19 pandemic and are not eligible for state or federal unemployment benefits because of their immigration status.
Monetary assistance			
13 (19.7)	5 (38.5)	8 (61.5)	AB 4171 (New Jersey): Would establish a one-time cash assistance program for eligible taxpayers. Would specify that all state taxpayers, including undocumented immigrants, are eligible to receive this stimulus payment. HB 3409 (Oregon): Would create a return-to-work incentive payment program for people who worked during the first year of the pandemic as frontline essential workers. Would specify that immigration status cannot be considered when deciding program eligibility. HB 138 (Vermont): Creates a state COVID-19 economic stimulus equity program. Offers eligibility to those who were ineligible to receive payment through the federal CARES Act because of their immigration status.

Note. CARES = Coronavirus Aid, Relief, and Economic Security.

[a]The “did not pass” number includes 5 bills that were introduced with language that addressed individuals who were undocumented within the context of the COVID-19 pandemic. However, this language was dropped from the bills as they worked their way through the legislative process. Versions of these bills—without language of relevance to the research question—were passed. Because of the removal of relevant language, these bills are categorized as “did not pass” for purposes of this study.

undocumented were eligible for these grants (California AB 132 [2021]). The final bills in this category were related to protections for undocumented workers who contracted COVID-19 while on the job (Rhode Island HB 5474 [2021]) and appropriating funds to conduct a study on frontline workers (Nevada SB 209 [2021]).

Monetary Assistance

Thirteen bills (19.7%) addressed the provision of monetary assistance during the pandemic. Of these, 5 (38.5%) were enacted into law. Eleven of the bills in this category sought to provide cash payments to individuals regardless of immigration status. For example, Washington State legislators passed into law a bill that provided stimulus payments to persons who were not eligible to receive federal economic impact payments because of their immigration status (Washington HB 1368 [2021]). The final 2 bills in this category would provide funds to cover COVID-19–related funeral expenses (California AB 868 [2021]) and offer grants for undocumented students (Washington SB 5451 [2021]).

DISCUSSION

Between January 2020 and August 2021, legislatures in 13 states introduced 66 bills and resolutions addressing people who were undocumented within the context of the COVID-19 pandemic. Although only 25.7% (17/66) ultimately became law, much can be learned from the legislation that was introduced. In particular, the 66 bills and resolutions demonstrate that state-level lawmakers were concerned about exclusion of people who were undocumented from the federal response to COVID-19 as well as the unique challenges faced by this population that may have been exacerbated by the pandemic.

Given the rise in anti-immigrant sentiment associated with the COVID-19 pandemic,[13,14] we considered whether legislation introduced at the state level would likely protect or harm people who were undocumented. Of the 66 bills and resolutions in our final data set, the vast majority (62/66; 94%) sought to protect or benefit this group. The remaining 4 bills and resolutions were neutral (e.g., creation of a task force). Given the uniquely vulnerable position of people who are undocumented[10,11]—and the fact that COVID-19 exacerbated many challenges this group was already facing[12]—it is not surprising that bills introduced during the pandemic generally sought to benefit this population. Moreover, these laws suggest a recognition that protecting the health of undocumented individuals was essential for combatting COVID-19 and ensuring economic recovery, thus benefiting the larger community.[22]

Our findings suggest that state-level lawmakers in some Democratic-controlled or split legislatures sought to fill perceived gaps in the federal pandemic response relative to those who were undocumented. Most notably, people who were undocumented were not eligible to receive the Economic Impact Payments (also known as “stimulus checks”) that were distributed by the federal government via the American Rescue Plan.[23] Some state legislatures sought to remedy this by introducing bills that would provide cash assistance to people regardless of their immigration status (e.g., California SB 86 [2021], New Jersey SB 2329 [2020], Washington HB 1368 [2021]). People who were undocumented were also excluded from temporarily expanded access to certain medical services during federally declared states of emergency. To address this, some state legislatures introduced bills that explicitly expanded health services (e.g., access to COVID-19 treatment and health plans’ essential health benefits) during declared emergency periods for people who were undocumented (e.g., Illinois SB 2294 [2021], New York SB 2549 [2021]).

State lawmakers were also sensitive to the unique circumstances potentially exacerbated by the pandemic for those who were undocumented. The most prominent example concerns bills that sought to prevent access to personal health information (e.g., contact tracing, vaccine status) by immigration authorities (e.g., Illinois HB 3120 [2021], Nevada AB 260 [2021], New York SB 900 [2021]). In addition to accounting for vulnerability among people who were undocumented, these bills demonstrate a recognition for the overriding importance of having complete and accurate health data to combat the pandemic.[24]

Several bills sought to address the especially vulnerable employment situation of undocumented laborers during COVID-19. For example, many direct services workers, who shouldered some of the highest-risk jobs during the pandemic, were individuals who were undocumented.[10] In recognition of this, some state legislatures sought to extend traditional forms of job protection and employment benefits—typically only available to US citizens and immigrants with documentation—to this group (e.g., Rhode Island HB 5474 [2021], Washington SB 5438 [2021]). In addition, some state legislatures introduced bills that included immigration status as a prohibited form of "retaliatory personnel action" to prevent employers from taking advantage of workers by threatening to report their undocumented status (e.g., Minnesota SF 1518 [2021], New Jersey SB 3827 [2021]).

Limitations

This research should be considered in light of several limitations. First, because we sought to provide a comprehensive exploration of legislative responses, our research did not capture other important state-level actions, such as executive orders or regulations. Second, our research scope did not include situating our findings within the context of legal responses to undocumented individuals during previous public health emergencies; however, this is a critical next step for future research.

Third, bills or resolutions may have been excluded if their text did not contain our search terms. To minimize the risk of this occurring, we created a comprehensive search string through an iterative process involving a literature review, extensive testing of the search string in Westlaw, and an expert review of the terminology. Fourth, although the research scope was to identify bills and resolutions that explicitly considered individuals who were undocumented within the context of the COVID-19 pandemic, we recognize that other proposed legislation may have had an impact on this population or had unintended consequences, even if it did not clearly focus on those who were undocumented.

Fifth, we recognize that analyzing the broader political context and implementation of passed legislation is critical for understanding the impact on undocumented individuals. However, this is outside the scope of our research, which sought to identify the introduction and passage of relevant bills and resolutions. Finally, relevant bills and resolutions could have been missed because they were introduced outside of our study's date range. However, given the trajectory of the pandemic, a representative sample of relevant legislation was likely introduced during our designated timeframe, which spans the prevaccine period, encompasses the duration of declared states of emergency, and captures the 18 months following WHO's March 2020 declaration of a global pandemic.

Public Health Implications

In response to the deadliest pandemic in US history, approximately one quarter of all state legislatures introduced bills or resolutions that addressed people who were undocumented within the context of COVID-19. Findings reveal how state-level legislators contemplated using the law to address the pandemic's disproportionate impact on an already highly vulnerable group. As states shift their attention from the exigencies of COVID-19, this research provides insight into how law might be used to protect those who are undocumented during the preparedness, response, and recovery stages of future public health emergencies. AJPH

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CONFLICTS OF INTEREST

The authors have no conflicts of interest to disclose.

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This project did not require institutional review board protocol approval because human participants were not involved in the research.

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Neighborhood Composition and Air Pollution in Chicago: Monitoring Inequities With a Dense, Low-Cost Sensing Network, 2021

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See also Shatas and Hubbell, p. 1693.

Objectives. To evaluate the efficacy of a novel, real-time sensor network for routine monitoring of racial and economic disparities in fine particulate matter ($PM_{2.5}$; particulate matter ≤ 2.5 μm in diameter) exposures at the neighborhood level.

Methods. We deployed a dense network of low-cost $PM_{2.5}$ sensors in Chicago, Illinois, to evaluate associations between neighborhood-level composition variables (percentage of Black residents, percentage of Hispanic/Latinx residents, and percentage of households below poverty) and interpolated $PM_{2.5}$. Relationships were assessed in spatial lag models after adjustment for all composition variables. Models were fit with data both from the overall period and during high-pollution episodes associated with social events (July 4, 2021) and wildfires (July 23, 2021).

Results. The spatial lag models showed that racial/ethnic composition variables were associated with higher $PM_{2.5}$ levels. Levels were notably higher in neighborhoods with larger compositions of Hispanic/Latinx residents across the entire study period and notably higher in neighborhoods with larger Black populations during the July 4 episode.

Conclusions. As a complement to sparse regulatory networks, dense, low-cost sensor networks can capture spatial variations during short-term air pollution episodes and enable monitoring of neighborhood-level inequities in air pollution exposures in real time. (*Am J Public Health*. 2022;112(12): 1765–1773. https://doi.org/10.2105/AJPH.2022.307068)

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Recognizing health equity as a priority, public health researchers and practitioners are increasingly seeking to monitor and mitigate disparities in exposures to preventable causes of disease.[1,2] Air pollution in the form of fine particulate matter ($PM_{2.5}$; particulate matter ≤ 2.5 μm in diameter) is a leading environmental contributor to disease burdens[3] and disparities in disease burdens.[4] Adding to the urgency is the role of climate change in increasing overall $PM_{2.5}$ exposures through longer and more extreme wildfire seasons,[5] although little is known about the extent to which these climate change–exacerbated pollution events affect inequities in exposures.

In the United States, public health agencies obtain $PM_{2.5}$ data from the Environmental Protection Agency (EPA) ambient air monitoring network. This network, implemented as a result of the Clean Air Act, has been credited with contributing to a reduction in $PM_{2.5}$ of approximately 70% since 1981.[6] However, remote sensing data offer evidence that areas with the highest air pollution exposures in 1981 remain the most polluted areas more than 30 years later.[6] Moreover, modeled estimates from EPA emissions inventories show that Black and Hispanic/Latinx people experience higher exposures than White people.[7] These findings provide a rationale for monitoring systems that track not only pollution exposures but also disparities in these exposures.

Existing EPA data provide accurate information on air pollution exposures but are subject to limitations related to data coverage over space and time. In Chicago, Illinois, a city of 600 square kilometers, the EPA maintains 4 stations monitoring $PM_{2.5}$. Real-time hourly estimates of $PM_{2.5}$ from monitors at 2 of these stations are available through the public AirNow tool,[8] which provides estimates of regional-level exposures; however, the data are too spatially sparse to allow inferences about more granular levels of exposure, which is a problem given that evidence from mobile monitoring campaigns shows air quality can vary significantly between neighborhoods and city blocks.[9] Alternatively, the EPA does provide a "downscaler model" of $PM_{2.5}$ estimates at the census tract level,[10] but the model cannot be used to track real-time exposures because the most recent data are from 2018.

Dense city-wide networks of low-cost sensors could complement existing regulatory networks, enabling routine real-time monitoring of spatial variations in environmental exposures. To date, however, cities in the United States have largely not designed and implemented their own dense sensing networks (with a few exceptions such as New York City[11]). Instead, cities monitoring air quality at "hyperlocal" levels rely on data from crowdsourced and mobile monitoring initiatives, approaches subject to limitations. Crowdsourced networks are poorly suited to monitoring disparities as a result of systematic biases in sensor locations; for example, commonly used PurpleAir sensors are more likely to be located in White areas and areas of high socioeconomic status than in areas with environmental justice concerns.[12,13]

Mobile monitoring campaigns, for which research-grade sensors are placed on moving vehicles, offer both city-wide coverage and insights on spatial variation[9]; if the vehicle fleet is small, however, this approach cannot compare multiple places at the same time or provide real-time insights for the city as a whole. Running larger campaigns would be prohibitively time and labor intensive for an urban public health department. Other common approaches (e.g., estimating $PM_{2.5}$ levels from satellite imagery, emissions inventories, or sophisticated chemical transport models) are subject to model-related uncertainties and would benefit from training and validation with additional data collected across diverse neighborhoods.[5] There is thus an opportunity for dense real-time monitoring to complement existing regulatory networks for the specific purpose of monitoring disparities in exposures.

In this study, we deployed a low-cost sensor network built to monitor racial/ethnic and economic inequities in air pollution exposures. We conducted our research in the city of Chicago, building on previous efforts to document heightened exposures in environmental justice neighborhoods[14,15] as well as evidence that heightened air pollution exposures and social vulnerabilities are clustered on the south and west sides of the city.[16] We evaluated data from July 2021 given that July historically has higher $PM_{2.5}$ readings in comparison with other months.[17] In addition, there were 2 air pollution episodes in July 2021: on July 4, an expected pollution episode contributed to large but short-term increases in $PM_{2.5}$, and on July 23, an unexpected pollution event that corresponded to wildfires similarly contributed to heightened $PM_{2.5}$ levels over a short period of time. We examined spatial clustering in $PM_{2.5}$ in relation to the spatial clustering of sociodemographic variables; we further evaluated relationships between neighborhood-level sociodemographic composition and $PM_{2.5}$.

METHODS

Chicago is a diverse city characterized by roughly equal thirds of White, Black, and Hispanic/Latinx populations. Chicago is also one of the most segregated cities in the United States and has seen concerns regarding structural racism as a fundamental cause of inequitable pollution burdens.[16]

Air Pollution

Network design. Our work relied on air pollution data from a novel network of 115 sensors located on bus shelters across Chicago. We deployed the network during the summer of 2021 in collaboration with the city of Chicago, the academic Array of Things initiative, and JCDecaux Chicago, the local affiliate of JCDecaux SA—the world's largest outdoor advertising company—which installed sensing devices on the city's bus shelters. We also collaborated with the Environmental Law and Policy Center to support neighborhood environmental justice organizations in reviewing the network design.

The network was designed with the aim of monitoring local inequities in air pollution exposures. For 80 devices, we selected sites using a stratified random sampling design based on the approach of the New York City Community Air Survey.[11] Of the remaining devices, 26 were allocated to sites selected by community partners, and 9 were sited across 3 EPA regulatory monitoring stations (3 per station). Additional details on the network design can be found in Appendix A (available as a supplement to the online version of this article at http://www.ajph.org). In this analysis, we used only sites allocated through stratified random sampling.

Eclipse device. We designed devices used in the network to provide real-time measurements of air pollution in an urban setting (for full details on the hardware, see Daepp et al.[18]). Each device included a Sensirion SPS30 sensor, which collected $PM_{2.5}$ readings every 5 minutes, as well as sensors for relative humidity, barometric pressure, and temperature. Further details on the Sensirion SPS30 are provided in Appendix A. In addition, details on our calibration function to improve sensor accuracy to levels consistent with EPA recommendations for low-cost sensors are provided in Appendix B (available as a supplement to the online version of this article at http://www.ajph.org). Of note, daily average sensor values were highly correlated with daily averages from regulatory monitors surrounding Chicago (for details, see Appendix B and Appendix C, Figure C1, available as a supplement to the online version of this article at http://www.ajph.org).

Data cleaning and processing. From July 2 to July 31, we obtained 568 156 5-minute readings from 78 sensors. Following Lu et al., we implemented a 4-step quality control procedure.[19] First, we removed sensor data that were deemed as malfunctioning, determined according to a moving 5-hour standard deviation of 0 (0% of readings). Second, we removed implausible readings by excluding values of 0 and values above the measurement range of 1000 micrograms per cubic meter per manufacturer specifications (0.003%). Third, using a 75% completeness criterion, we removed readings from hours with less than 9 (of 12) 5-minute measures and from days with less than 18 (of 24) hours of data (1.86%). Finally, as a secondary check for malfunctioning devices, we assessed the extent to which sensor readings were consistent with readings from neighboring sensors.

In addition, we performed a linear regression of daily average readings for a given index sensor and its neighbors within a 5-kilometer radius and removed all readings from the index sensor if the R^2 value was less than 0.6 (0.01% of readings, affecting only a single sensor). With these criteria, 557 457 (98.1%) of the original 5-minute readings remained.

Daily average $PM_{2.5}$ values were calculated for each sensor by initially averaging 5-minute readings for each hour within a specified day (up to 24 hours), calibrating hourly data, and then aggregating those hourly averages for each specified day. With the exclusion criteria, 77 sensors remained. Preliminary analyses assessing spatial clustering in the set of device days excluded suggested no substantial clustering (Appendix C, Figure C2, available as a supplement to the online version of this article at http://www.ajph.org).

Spatial interpolation. We used inverse distance weighting (IDW) to estimate $PM_{2.5}$ across 77 community areas for the entire study period and during 2 air pollution episodes (July 4, consistent with excess air pollution caused by fireworks,[20] and July 23, consistent with nationwide increases in air pollution caused by wildfires on the west coast[21]). Chicago is divided into 77 community areas covering an average of 7.8 square kilometers. These community areas have historically been used for planning and statistical purposes and remain largely consistent with residents' contemporary perceptions of neighborhood boundaries.[22]

IDW interpolation allowed us to predict $PM_{2.5}$ values across unknown points (i.e., reference points) on the basis of nearby points where $PM_{2.5}$ is known (i.e., monitoring points). The approach assigns values to reference points through a weighted average of the values at monitoring points; monitoring points closest to a given reference point have larger weights than monitoring points further away. Weights are defined by the inverse of the distance between each reference point and monitoring point and then raised to an arbitrary power that we set equal to 2 (i.e., the square of the inverse distance), a value supported by both empirical cross validation and previous literature.[23]

We used IDW to create smoothed maps (rasters) of averaged estimated $PM_{2.5}$ at a grid cell resolution of 100 × 100 meters. We then aggregated these grid cells to compute community area–level averages such that a grid cell belonged to a given community area if its centroid fell inside of it. We further evaluated the robustness of our results to the use of a different interpolation approach, ordinary kriging; the 2 approaches produced similar neighborhood-level estimates (Spearman's ρ = 0.86), and thus we used the IDW approach because of its interpretability and consistency with theoretical models of air pollution spread[23] and its widespread usage both in academic research[23,24] and by the EPA.[25]

Sociodemographic Composition Variables

We used sociodemographic data from the 2015 to 2019 American Community

Survey[26] to measure racial/ethnic composition and the percentage of households below poverty. Regarding racial/ethnic composition, we focused on percentages of non-Hispanic Black and Hispanic/Latinx residents because these are the 2 largest racial/ethnic minority groups in Chicago, accounting for 29.1% and 28.7% of the city's population, respectively. Census tract–level measures were aggregated to community areas such that a census tract belonged to a given community area if its centroid fell inside of it.

Statistical Analysis

Initially, we computed descriptive statistics for each outcome (interpolated $PM_{2.5}$ values for July 2–31, July 4, and July 23) and composition variable. We also assessed the degree of spatial autocorrelation for each variable by calculating global Moran's *I* statistics. Moran's *I* values range from −1 to 1; values near 0 indicate no autocorrelation (i.e., randomness), positive values indicate clustering, and negative values indicate dispersion. We then mapped the distribution of $PM_{2.5}$ and each composition variable at the community area level. In addition, we assessed bivariate relationships between tract-level composition variables and $PM_{2.5}$ using Spearman correlation coefficients.

To model the relationship between sociodemographic composition and $PM_{2.5}$, we created linear regression models adjusting for each socioeconomic composition measure over the entire study period and separately for the July 4 and July 23 air pollution episodes. We adjusted for all 3 measures to avoid overestimating the effects of any single variable. Because air pollution is spatially patterned, violating the linear regression assumption of independence,[27] we also fit a series of spatial lag models to account for spatial dependence.

Spatial lag models are similar to linear regression models, but they include in addition a lagged dependent variable reflecting the weighted average of $PM_{2.5}$ values across neighboring community areas. The coefficient associated with this lagged variable (ρ) quantifies the strength of spatial dependence. If ρ is greater than 0, this indicates that $PM_{2.5}$ values are positively related to those of neighboring community areas; a negative value indicates the inverse. A value of 0 indicates no dependence and renders the equation equivalent to a linear model.[28] Spatial lag models, relative to other spatial regression models such as spatial error models, are often used when researchers believe that spatial autocorrelation is caused by an underlying substantive process.[28,29] In the case of air pollution, we theorized that industrial zoning and related planning policies that resulted in clusters of pollution sources were ultimately caused by distal, structural processes of discrimination and residential segregation.[30]

To evaluate the presence of autocorrelation in our regression models, we calculated Moran's *I* values for residuals using first-order queen contiguity-based weights. Models incorporating queen contiguity-based weights yielded lower Akaike information criterion (AIC) values than preliminary models employing (1) rook contiguity-based weights and (2) the minimum distance for all community areas to have at least 1 neighbor and were thus chosen for our analysis. Pseudo *P* values for Moran's *I* statistics were generated via a Monte Carlo simulation of 999 random replications. We considered autocorrelation to be present if pseudo *P* values were less than .05.

We used R version 4.1.0 (R Foundation for Statistical Computing, Vienna, Austria) in conducting our analyses; we used the *gstat* package to perform IDW and the *spdep* package to generate spatial lag models. All models were fit with community area–level data; as a sensitivity analysis addressing concerns regarding the modifiable areal unit problem, we replicated our main analyses at the census tract level.

RESULTS

Table 1 summarizes descriptive statistics across the 77 community areas. The average level of interpolated $PM_{2.5}$

TABLE 1— Descriptive Statistics of Variables Across 77 Community Areas: Chicago IL, July 2–31, 2021

	Mean (SD)	Range
Interpolated $PM_{2.5}$ (µg/m^3)		
July 2–31	13.2 (0.7)	11.3–14.6
July 4	14.2 (1.5)	10.6–17.3
July 23	26.6 (1.0)	24.1–28.3
Percentage		
Black	38.1 (39.1)	0.4–96.5
Hispanic/Latinx	26.0 (26.9)	0.1–89.2
Households below poverty	19.6 (10.9)	3.5–53.9
Eclipse devices, no.	1.0 (0.9)	0.0–4.0

Note. $PM_{2.5}$ = particulate matter ≤ 2.5 µm in diameter.

during the study period was 13.16 μg/m^3 (range = 11.3–14.6 μg/m^3). Average interpolated $PM_{2.5}$ levels were slightly elevated on July 4 (14.2 μg/m^3; range = 10.6–17.3 μg/m^3) and doubled on July 23 (26.6 μg/m^3; range = 24.1–28.3 μg/m^3). The network design resulted in an average of 1 sensor allocated to each community area (range = 0–4).

The average community area was composed of 38.1% Black residents (range = 0.4%–96.5%) and 26.0% Hispanic/Latinx residents, with an average of 19.6% of households below poverty (range = 3.5%–53.6%). Consistent with known racial and economic segregation patterns in Chicago, as shown in Appendix C, Figure C3 (available as a supplement to the online version of this article at http://www.ajph.org), Black residents were clustered in areas on the west and south sides (Moran's I = 0.71; P = .001); Hispanic/Latinx residents were clustered in areas on the northwest, southwest, and south sides (Moran's I = 0.63; P = .001); and the percentage of households below poverty was generally larger along the outer edge of the city (Moran's I = 0.51; P = .001).

Figure 1 illustrates the spatial distribution of interpolated $PM_{2.5}$ values during the entire study period and for each air pollution episode. Interpolated values of $PM_{2.5}$ displayed substantial spatial clustering, indicated by large Moran's I values (overall: 0.74; July 4: 0.81; July 23: 0.66; all pseudo Ps = .001). Notably, higher $PM_{2.5}$ levels appeared to cluster primarily along the west side during the study period overall and the July 23 pollution episode, whereas higher levels appeared to cluster along the south side during the July 4 pollution episode.

We provide aspatial Spearman correlation coefficients in Appendix D, Table D1 (available as a supplement to

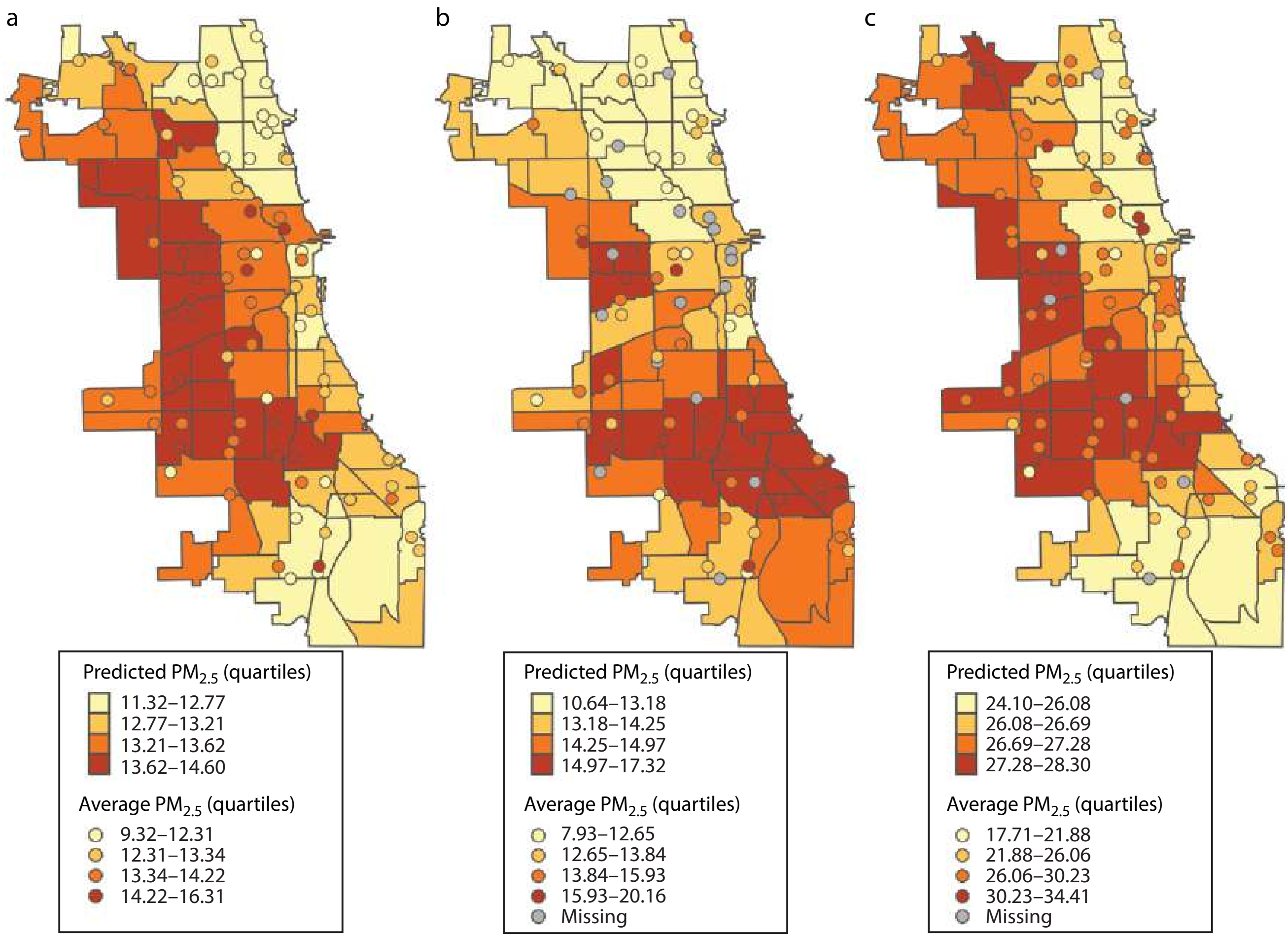

FIGURE 1— Inverse Distance Weighted Predicted $PM_{2.5}$ Levels Summarized to Community Areas and Average Values From Project Eclipse Sensors in Chicago, IL: (a) July 2–31, 2021; (b) July 4, 2021; and (c) July 23, 2021

Notes. $PM_{2.5}$ = fine particulate matter ≤ 2.5 μg/m^3. There were 77 community areas and 77 sensors. Moran's I = 0.74, 0.81, and 0.66 for parts a, b, and c, respectively.

the online version of this article at http://www.ajph.org), and focus the remainder of our results on regression models. Specifically, we focus on adjusted spatial lag models, as Moran's *I* values were reduced to nearly 0 and pseudo *P* values were all above .05, indicating that spatial models sufficiently removed spatial autocorrelation; furthermore, these models achieved better fits than corresponding linear regression models. For the period from July 2 to July 31, Table 2 shows substantial evidence of a positive relationship between Hispanic/Latinx residential composition and $PM_{2.5}$ (B = 0.47; 95% confidence interval [CI] = 0.10, 0.84). Relationships between Black residential composition and percentage of households below poverty were also positive but marked with imprecision.

During the July 4 episode, the adjusted spatial model suggested positive, substantial relationships between $PM_{2.5}$ and both Black residential composition (B = 1.13; 95% CI = 0.44, 1.82) and Hispanic/Latinx residential composition (B = 0.67; 95% CI = 0.01, 1.33). The relationship between percentage of households below poverty and $PM_{2.5}$ was positive but marked with imprecision once again. Finally, during the July 23 episode, the adjusted spatial model showed that all sociodemographic composition variables had a positive relationship with $PM_{2.5}$ but were marked with imprecision.

Sensitivity analyses in which census tracts were used yielded results similar

TABLE 2— Adjusted Models of Sociodemographic Composition on $PM_{2.5}$ Across 77 Community Areas: Chicago, IL, July 2–31, 2021

	Linear Model	Spatial Lag Model
July 2–31		
% Black, B (95% CI)	0.55 (−0.09, 1.18)	0.12 (−0.21, 0.45)
% Latinx, B (95% CI)	1.61 (0.93, 2.30)	0.47 (0.10, 0.84)
% of households below poverty, B (95% CI)	1.13 (−0.74, 2.99)	0.79 (−0.15, 1.74)
ρ, B (95% CI)	...	0.85 (0.75, 0.95)
Moran's *I* (*P*)	0.63 (.001)	0.01 (.33[a])
AIC	148.63	70.65
Log likelihood (*df*)	−69.32 (5)	−29.32 (6)
R^2	0.29	0.75[b]
July 4		
% Black, B (95% CI)	3.69 (2.73, 4.66)	1.13 (0.44, 1.82)
% Latinx, B (95% CI)	2.65 (1.61, 3.69)	0.67 (0.01, 1.33)
% of households below poverty, B (95% CI)	−0.14 (−2.98, 2.70)	0.24 (−1.46, 1.95)
ρ, B (95% CI)	...	0.76 (0.63, 0.88)
Moran's *I* (*P*)	0.39 (.001)	0.01 (.4[a])
AIC	213.47	155.76
Log likelihood (*df*)	−101.74 (5)	−71.88 (6)
R^2	0.61	0.82[b]
July 23		
% Black, B (95% CI)	0.17 (−0.82, 1.15)	0.13 (−0.42, 0.67)
% Latinx, B (95% CI)	1.20 (0.14, 2.27)	0.35 (−0.24, 0.94)
% of households below poverty, B (95% CI)	1.61 (−1.30, 4.51)	0.41 (−1.18, 1.99)
ρ, B (95% CI)	...	0.86 (0.75, 0.96)
Moran's *I* (*P*)	0.61 (.001)	0.09 (.1[a])
AIC	217.19	150.45
Log likelihood (*df*)	−103.60 (5)	−69.22 (6)
R^2	0.10	0.63[b]

Note. AIC = Akaike information criterion; CI = confidence interval; $PM_{2.5}$ = particulate matter ≤ 2.5 μm in diameter. The sample size was 77.

[a]Pseudo *P* value.

[b]Nagelkerke pseudo R^2 value.

to those of our main analyses at the community area level, albeit with increased precision and substantial autocorrelation in adjusted spatial lag models (Appendix D, Table D2, available as a supplement to the online version of this article at http://www.ajph.org).

DISCUSSION

In this study, we examined differences in sociodemographic disparities in $PM_{2.5}$ across long-term monitoring periods versus during short-term air pollution episodes. We have provided evidence of spatial variation across neighborhoods in levels of $PM_{2.5}$ both across the study period and during specific high-pollution episodes characterized by social events (July 4) and wildfires (July 23). Exposures were substantially higher in areas with larger compositions of Hispanic/Latinx residents across the entire study period and substantially higher in areas with larger compositions of Hispanic/Latinx and Black residents during the July 4 episode. No sociodemographic composition variable was associated with $PM_{2.5}$ during the July 23 episode. Our results demonstrate the effectiveness of a city-wide, real-time sensing network for measuring ongoing and episodic neighborhood-level disparities in poor air exposures.

Evidence of heightened $PM_{2.5}$ in areas with relatively more Hispanic/Latinx and Black residents is consistent with literature documenting racial and ethnic disparities in $PM_{2.5}$[31] as well as studies linking racial residential segregation to environmental disparities.[30,32] Our study adds to this body of literature in 2 key ways. First, we demonstrated differences in the groups affected in short-term episodes versus over longer time periods. Although neighborhoods with larger proportions of Hispanic/Latinx residents appeared to have the largest $PM_{2.5}$ burden overall, July 4 may be an especially harmful pollution event disproportionately affecting areas with larger proportions of Black residents. These results point to the need for more targeted interventions that consider both spatial and temporal contexts. Although similar findings might be obtained by downscaling models that combine regulatory monitoring data with chemical transport models or satellite data, a particular benefit of a real-time monitoring approach is its potential to make findings available to policymakers immediately after or even during a pollution episode, supporting mitigation efforts.

Second, our work offers evidence that city-wide high-pollution events such as the wildfire-related episode on July 23 may result in minimal variations across sociodemographic composition, making dense, real-time monitoring particularly beneficial if such events obscure disparities occurring during more typical $PM_{2.5}$ exposure days.

Limitations

Our study is subject to several key limitations. First, logistical delays in sensor deployment resulted in missing data over space and time. Of the 80 sensors intended for analysis in this study, 3 could not be deployed until the end of July; a slow rollout over the first week of monitoring as well as occasional intermittent sensor failures further limited the sample size to 59 sensors on July 4 and 71 on July 23. However, the data collected were robust against data quality issues, as our quality control procedure identified issues with less than 2% of the 5-minute readings collected and no spatial clustering was detected in the fraction of missing device days.

Second, low-cost sensors can exhibit low accuracy for regulatory purposes. We sought to address this limitation through calibration and focusing on the comparison of trends over time and of sensors with each other; however, exact estimates should be treated with caution because the results may have been affected by systematic sensor error. Our network thus cannot substitute for regulatory networks; rather, as a complement to sparse regulatory monitors, our approach can help prioritize monitoring and mitigation of disparities in air pollution. We were also able to reduce random error by aggregating data to daily average values; however, this methodological decision limited our ability to take advantage of the high temporal resolution of the sensor readings.

Third, the generalizability of our results may be limited by our siting of sensors at bus shelters. However, there were also benefits to the use of bus shelters; all devices were placed in similar contexts and at a consistent height, and, because technicians regularly visit these bus shelters, the siting approach facilitated maintenance that mitigated data loss.[18] Moreover, bus shelters represent locations where people congregate and—importantly—breathe. Also, our analyses were limited to July 2021 and may not generalize to other months; however, our work does provide a framework for extending such a monitoring network over longer periods.

Finally, land use regression models could offer increased precision relative to IDW in estimating hyperlocal variations in $PM_{2.5}$. However, these models incorporate temporally invariant covariates (e.g., physical geography) that would smooth out key real-time fluctuations. A lack of ground-truth data beyond what the 4 EPA sites capture further limited our ability to validate our models.

Policy Implications

We have described the development of a dense, real-time sensor network that enables characterization of spatial variations in $PM_{2.5}$ at the neighborhood scale. Examining a period characterized by multiple short-term air pollution episodes, we showed persistent inequities throughout the study period as well as important variations in the groups most affected by different short-term events. It follows that interventions seeking to address inequities in air pollution exposures may need to contend with how inequities vary across time and specific pollution events. Our results show how low-cost sensors can be used in a large, urban setting for monitoring environmental inequities, offering an approach that can be reproduced by public health departments in other cities seeking to promote environmental justice. AJPH

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CONTRIBUTORS

P. Esie, M. I. G. Daepp, and S. Counts contributed to study design and writing. M. I. G. Daepp, A. Roseway, and S. Counts contributed to data collection. P. Esie and M. I. G. Daepp contributed to data analysis. All of the authors contributed to study conception.

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CONFLICTS OF INTEREST

The authors declare no conflicts of interest.

HUMAN PARTICIPANT PROTECTION

Because no human participants were involved, this study was exempt from institutional board review.

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Living Alone and Suicide Risk in the United States, 2008–2019

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See also Shaw, p. 1699 and Nestadt, p. 1702.

Objectives. To evaluate the association between living alone and suicide and how it varies across sociodemographic characteristics.

Methods. A nationally representative sample of adults from the 2008 American Community Survey (n = 3 310 000) was followed through 2019 for mortality. Cox models estimated hazard ratios of suicide across living arrangements (living alone or with others) at the time of the survey. Total and sociodemographically stratified models compared hazards of suicide of people living alone to people living with others.

Results. Annual suicide rates per 100 000 person-years were 23.0 among adults living alone and 13.2 among adults living with others. The age-, sex-, and race/ethnicity-adjusted hazard ratio of suicide for living alone was 1.75 (95% confidence interval = 1.64, 1.87). Adjusted hazards of suicide associated with living alone varied across sociodemographic groups and were highest for adults with 4-year college degrees and annual incomes greater than $125 000 and lowest for Black individuals.

Conclusions. Living alone is a risk marker for suicide with the strongest associations for adults with the highest levels of income and education. Because these associations were not controlled for psychiatric disorders, they should be interpreted as noncausal. (*Am J Public Health*. 2022;112(12):1774–1782. https://doi.org/10.2105/AJPH.2022.307080)

AJPH December 2022, Vol 112, No. 12

Between 1960 and 2021, the percentage of single-person households in the United States increased from 13% to 28%.[1] One-person households also account for more than a quarter of all households in many other high-income countries including France, England, Germany, Canada, Spain, and Japan.[2] In light of the substantial number and rising proportion of adults who live alone, there is interest in understanding whether and to what extent living alone is associated with adverse health outcomes.

Several general population cohort studies have reported that living alone is connected with increased risk of all-cause mortality. In one review, the average increased risk of all-cause mortality for living alone (32%) was similar to the corresponding risks for social isolation (29%) and loneliness (26%).[3] A recent meta-analysis reported that living alone is associated with increased risk of all-cause mortality for individuals aged younger than 65 years and may be more pronounced for males than females.[4] Informed by social and psychological theories linking social isolation to suicide risk,[5] several studies have specifically probed relationships between living alone and risk of suicide. Cohort studies of various high-risk populations including adults following nonfatal suicide attempts,[6] people with disabilities attributable to mental disorders,[7] adults with bipolar disorder,[8] and people hospitalized for depression[9] have all reported significant positive associations between living alone and suicide risk.

In general population samples, living alone has also been reported to be associated with increased risk of suicide. A German population-based cohort study reported that living alone was associated with increased risk of suicide (hazard ratio [HR] = 2.2) similar in magnitude to depressed mood (HR = 2.0).[10] A large Finnish general population cohort study further reported that living alone was associated with increased relative suicide mortality rates for men and women who were working age (30–64 years) and older (≥ 65 years).[11] A recent UK Biobank study, however, found that living alone was associated with an increased risk of suicide in men but not women.[12] A study of older Korean adults that controlled for a wide range of sociodemographic, health, and behavioral health factors similarly found that living alone was related to suicidal ideation for men but not women.[13] Some[14,15] but not all[16] case–control studies have also reported significant associations between living alone and death by suicide.

Because of sample size limitations of previous research, little is known about whether and how the risk of suicide associated with living alone varies across sociodemographic groups beyond the apparent stronger association for men than women. The multiple pathways to living alone, which include relationship dissolution, death of a partner, and decisions not to enter into a cohabitation partnership, contribute to the heterogeneity of this population, and the mental health consequences of living alone could vary across this diverse group.

To better understand the association between living alone and suicide, we followed respondents to the 2008 American Community Survey (ACS) who were either living alone or with others for their risk of death by suicide. Stratified analyses assessed whether living alone varied as a risk marker for suicide across sociodemographic groups. Because the ACS does not include measures of common shared causes of living alone and suicide, such as mental health problems[17,18] and substance misuse,[19,20] we consider these associations as noncausal. Increasing our understanding of the strength and pattern of associations between living alone and suicide might inform risk assessment and future epidemiological research to evaluate the contribution of living alone to suicide risk.

METHODS

The study cohort was defined from the Mortality Disparities in American Communities[21,22] sample that links 2008 ACS data to National Death Index underlying cause of death certificate records from 2008 to 2019 (n = 3 452 000) after exclusion of people for whom National Death Index linkage was not possible because social security numbers, names, and date of birth were unavailable. The complex sampling frame of the ACS was designed to approximate US population estimates by age, sex, race/ethnicity, and state of residence. Sampling weights were applied to account for variable sampling within demographic subgroups.

We analyzed respondents aged 18 years or older at the ACS interview, excluding those living in group quarters (n = 142 000) such as college dormitories, residential treatment centers, skilled nursing facilities, group homes, military barracks, or correctional facilities.

Living Alone

The number of persons in a household was defined as everyone currently living or staying at a sampled address, except those who have been or will be living at the address for 2 months or less. The study cohort was partitioned into 2 groups on the basis of their reported living circumstances: (1) adults living alone or (2) adults living with others including family and nonfamily. The living-alone variable was measured once in 2008.

Sociodemographics and Functional Disabilities

Respondent characteristics were collected at the time of the ACS survey. Sociodemographic characteristics included age in years, sex, race/ethnicity, marital status, employment during past week, highest level of educational attainment, household annual income from all sources, urban (77%) or rural (23%) residence as defined by the Census,[23] whether the respondent was a renter or owner, and residential stability based on how long the respondent had lived at their current residence (< 5 years, 5–10 years, > 10 years).

Respondents were also asked about 6 areas of serious difficulties including hearing; vision; concentrating, remembering, or making decisions; walking or climbing stairs; dressing or bathing; and independent living. Respondents who indicated 1 or more of these difficulties were coded as having "any functional disability."

Outcome

National Death Index data indicated whether each Mortality Disparities in American Communities participant had died over the 11-year follow-up period from their ACS survey date. The outcome of primary interest was suicide (*International Classification of Diseases and Related Health Problems, 10th Revision, Clinical Modification* [*ICD-10-CM*; Second edition; Geneva, Switzerland: World Health Organization; 2004] codes X60–X84, Y87.0, U03)[24] as the underlying cause of death.

Statistical Analysis

The analysis was performed in 3 stages. In the first stage, we used the χ^2 difference in proportion test to compare the sociodemographic characteristics of adults who lived alone versus with others. In the second stage, we determined suicide rates per 100 000 person-years with 95% confidence intervals (CIs). We also examined whether each sociodemographic characteristic moderated the strength of living alone as a risk marker for suicide. Because living alone[25] and suicide[26] both vary by age, sex, and race/ethnicity, we also treated these demographic characteristics as potential background confounders. Therefore, we used Cox proportional hazards models, adjusted for age, sex, and race/ethnicity, to estimate adjusted hazard ratios (AHRs) of suicide with living alone as the independent variable of interest and living with others as the reference group.

We measured event time continuously from the date of baseline survey administration until the date of suicide death, date of death from all causes other than suicide (censoring event), or December 31, 2019, for those who did not die (censoring event), whichever came first. A survival plot was generated to display cumulative suicide risks for respondents living alone and with others. In separate models, we entered interaction terms (e.g., age group × living situation) to test whether the effects of living situation on hazards of suicide differed across levels of the sociodemographic variables. Separate analyses partitioned suicide deaths by means into poisoning (*ICD-10-CM*: X60–X69), firearms (X72–X74), suffocation (X70), and other (X71, X75–X84, Y87.0, U03).

In a sensitivity analysis, we limited follow-up to 1 year from ACS completion. In a second sensitivity analysis, we broadened the definition of mortality outcome to include suicide (*ICD-10-CM*: X60–X84, Y87.0, U03) or injuries of undetermined intent (Y10–Y34, Y87.2). We considered rates and AHRs with nonoverlapping 95% CIs or *P* value less than .05 to significantly differ.

We conducted analyses in SAS version 9.4 (SAS Institute, Cary, NC). We weighted individual-level observations to account for nonequal probability of selection into ACS and to increase generalizability of the findings to the US adult population. Reporting followed the disclosure guidelines of the Census Bureau's Disclosure Review Board.

RESULTS

Approximately 14.5% of the sample, including 16.3% of women and 12.6% of men, lived alone at the time of the survey. As compared with people who lived with others, those who lived alone were significantly older and were more likely to be female, to have White or Black race/ethnicity, to have a low income, to reside in more urban rather than the most rural areas, to rent rather than own their residence, and to have a functional disability. However, people who lived alone were less likely than those who lived with others to be employed or to be currently married (Table 1).

Overall and Stratified Risk of Suicide

The overall annual rate of suicide per 100 000 person-years was nearly twice as high among people who lived alone compared with people living with others (23.0 vs 13.2; Table 2). Group differences in the cumulative risk of suicide during follow-up are displayed in Figure 1 (Wald $\chi^2 = 268.3$; $P < .001$). After we controlled for the potentially confounding effects of age, sex, and race/ethnicity, living alone was also associated with nearly 2-fold increased hazards of suicide in the total sample (AHR = 1.75; 95% CI = 1.64, 1.87). Across most strata examined, adults who lived alone had significantly higher hazards of suicide than people who lived with others. The 2 strongest associations of living alone with suicide risk were among adults with a bachelor's degree or higher education (AHR = 2.25; 95% CI = 1.97, 2.56) and among adults with annual incomes of more than $125 000 (AHR = 2.22; 95% CI = 1.64, 3.00) while the 2 weakest corresponding associations were among non-Hispanic Black adults (AHR = 0.92; 95% CI = 0.63, 1.33) and among adults aged 18 to 39 years (AHR = 1.23; 95% CI = 1.07, 1.41).

We observed significant variations in the adjusted hazards of suicide risk by age group, sex, race/ethnicity, education, income, and functional disability status (Table 2). Specifically, the association between living alone and suicide was significantly stronger for older (AHR = 1.97; 95% CI = 1.68, 2.31) than younger (AHR = 1.23; 95% CI = 1.07, 1.41) adults, men (AHR = 1.82; 95% CI = 1.69, 1.97) than women (AHR = 1.69; 95% CI = 1.45, 1.97), non-Hispanic White (AHR = 1.79; 95% CI = 1.67, 1.93) than non-Hispanic Black (AHR = 0.92; 95% CI = 0.63, 1.33) individuals, and people with a bachelor's degree or higher education (AHR = 2.25; 95% CI = 1.97, 2.56) than for those whose with less than a high-school education (AHR = 1.77; 95% CI = 1.46, 2.16).

The association between living alone and suicide hazards was also stronger for people whose annual incomes exceeded $125 000 (AHR = 2.22; 95% CI = 1.64, 3.00) than for those with incomes below $40 000 (AHR = 1.38; 95% CI = 1.26,1.52). In addition, living alone was associated with significantly greater hazards of suicide for people living without functional disabilities (AHR = 1.77; 95% CI = 1.64, 1.91) than for those living with these disabilities (AHR = 1.49; 95% CI = 1.31, 1.71) as was the associations with owners (AHR = 1.83; 95% CI = 1.67, 2.00) than renters (AHR = 1.56; 95% CI = 1.40, 1.74). In sex-stratified analyses, there were several similarities between the

TABLE 1— Sociodemographic Characteristics of Adults Who Live Alone or With Others: Mortality Disparities in American Communities, United States, 2008

Characteristic	Adults Living Alone (n = 480 000), % (95% CI)[a]	Adults Living With Others (n = 2 830 000), % (95% CI)[a]
Age, y*		
18–39	22.9 (22.7, 23.0)	41.8 (41.7, 41.9)
40–64	43.4 (43.2, 43.5)	44.4 (44.4, 44.5)
≥ 65	33.8 (33.6, 34.0)	13.8 (13.7, 13.8)
Sex*		
Male	44.2 (44.0, 44.3)	49.0 (48.9, 49.1)
Female	55.8 (55.7, 56.0)	51.0 (50.9, 51.1)
Race/ethnicity		
Non-Hispanic White	74.9 (74.8, 75.1)	67.8 (67.7, 67.9)
Non-Hispanic Black	13.6 (13.5, 13.8)	10.8 (10.8, 10.9)
Hispanic	7.1 (7.0, 7.2)	14.6 (14.5, 14.6)
Other	4.4 (4.3, 4.5)	6.8 (6.8, 6.8)
Marital status*		
Married	4.1 (4.0, 4.2)	62.4 (62.3, 62.4)
Separated/divorced	35.0 (34.8, 35.1)	10.0 (9.9, 10.0)
Widowed	25.6 (25.5, 25.8)	3.1 (3.0, 3.1)
Never married	35.3 (35.2, 35.5)	24.6 (24.5, 24.7)
Employment*		
Employed	55.7 (55.6, 55.9)	66.4 (66.3, 66.5)
Not employed, < 65 y	14.9 (14.8, 15.0)	22.1 (22.1, 22.2)
Not employed, ≥ 65 y	29.4 (29.2, 29.5)	11.5 (11.4, 11.5)
Education*		
Less than high school	14.5 (14.3, 14.6)	15.0 (14.9, 15.0)
High school/GED	27.5 (27.4, 27.7)	28.7 (28.7, 28.8)
Some college/associate degree	29.3 (29.2, 29.5)	30.8 (30.7, 30.8)
Bachelor's degree or higher	28.7 (28.5, 28.8)	25.5 (25.5, 25.6)
Income,* $		
0 to 40 000 (loss)	67.1 (67.0, 67.3)	26.2 (26.1, 26.2)
40 001 to 75 000	22.0 (21.8, 22.1)	30.1 (30.0, 30.2)
75 001 to 125 000	7.8 (7.7, 7.8)	26.1 (26.0, 26.1)
> 125 000	3.2 (3.1, 3.2)	17.7 (17.6, 17.7)
Residence*		
Urban	82.1 (81.9, 82.2)	75.7 (75.6, 75.7)
Rural	17.9 (17.8, 18.1)	24.3 (24.3, 24.4)
Housing finance*		
Renter	46.2 (46.0, 46.4)	72.8 (72.7, 72.8)
Owner	53.8 (53.6, 54.0)	27.2 (27.2, 27.3)
Residential stability,* y		
< 5	44.8 (44.6, 45.0)	40.2 (40.2, 40.3)
5–10	19.8 (19.7, 20.0)	23.1 (23.0, 23.2)
> 10	35.4 (35.2, 35.5)	36.7 (36.6, 36.7)

Continued

associations among men and women (Tables A and B, available as supplements to the online version of this article at https://ajph.org). Among Hispanic adults, however, there was a significant association between living alone and suicide for men (AHR = 2.51; 95% CI = 1.84, 3.43) but not for women (AHR = 1.00; 95% CI = 0.36, 2.76).

In an analysis limited to 1-year follow-up after ACS completion, living alone was associated with increased hazards of suicide (AHR = 1.68; 95% CI = 1.39, 2.03; Table C, available as a supplement to the online version of this article at https://ajph.org) that were similar to the increase after the 11-year follow-up (AHR = 1.75; 95% CI = 1.64, 1.87; Table 2).

Risk of Suicide by Different Means

The adjusted hazards of suicide of living alone compared with living with others were higher for suicide by poisoning (AHR = 2.29; 95% CI = 1.97, 2.68) than by firearms (AHR = 1.69; 95% CI = 1.54, 1.85), suffocation (AHR = 1.52, 95% CI = 1.29, 1.78), or other means (AHR = 1.75; 95% CI = 1.64, 1.88; Table D, available as a supplement to the online version of this article at https://ajph.org).

Risk of Undetermined Intent Deaths

Broadening the outcome to suicide or undetermined intent injury deaths yielded rates per 100 000 person-years of 25.3 for adults living alone and 14.7 for adults living with others with an AHR of 1.74 (95% CI = 1.63, 1.85; Table E, available as a supplement to the online version of this article at https://ajph.org). The pattern of results with this broader

AJPH December 2022, Vol 112, No. 12

TABLE 1— Continued

Characteristic	Adults Living Alone (n = 480 000), % (95% CI)[a]	Adults Living With Others (n = 2 830 000), % (95% CI)[a]
Functional disability		
Any*	25.1 (24.9, 25.2)	12.8 (12.8, 12.9)
Hearing*	7.8 (7.7, 7.9)	3.8 (3.8, 3.9)
Vision*	5.1 (5.0, 5.2)	2.4 (2.4, 2.4)
Cognitive*	7.6 (7.6, 7.7)	4.4 (4.3, 4.4)
Walking*	15.9 (15.8, 16.0)	7.0 (7.0, 7.1)
Dressing*	4.9 (4.8, 5.0)	2.6 (2.5, 2.6)
Independent travel*	9.9 (9.8, 10.0)	4.8 (4.8, 4.9)

Notes. CI = confidence interval; GED = general educational development. Limited to adults aged ≥ 18 years; excludes adults in group quarters.

[a]Numbers rounded to 10 000s following Census guidelines. Disclosure Review Board approval number CBDRB-FY22-CES004-040.
*$P < .001$.

outcome resembled the pattern with suicide as the outcome (Table 2).

DISCUSSION

In this large, nationally representative cohort of US adults, living alone emerged as a significant risk marker for suicide. The strength of the association in the total adult population, which increased by 75% the hazards of suicide after controlling for age, sex, and race/ethnicity, was in line with previous epidemiological research from outside the United States.[10–12] Living alone was significantly associated with suicide mortality separately for men and women. There was significant variation across sociodemographic groups in the adjusted strength of the associations between living alone and suicide with the 2 strongest associations occurring among adults with the highest levels of income and education.

Because the present study did not control for psychiatric morbidity or substance use, which are related to living alone and suicide, the associations should be interpreted as noncausal. However, previous research on this topic, which controlled for different aspects of psychiatric morbidity or substance use, suggests living alone contributes to suicide risk. In a general population study, which controlled for baseline depressed mood, alcohol intake, and several other factors, living alone was associated with increased suicide risk (HR = 2.19; 95% CI = 1.09, 4.37).[10] A case–control study that matched on background demographic characteristics and controlled for psychiatric pathology further reported a significant association between living alone and suicide (odds ratio = 2.30; 95% CI = 1.36, 5.75).[14] Significant associations between living alone and suicide have also been reported in cohort studies restricted to individuals with psychiatric disorders[7–9] or following nonfatal intentional poisonings.[6]

Comparing the background characteristics of adults who lived either alone or with others suggests that living alone is related to a set of socioeconomic and functional vulnerabilities. In relation to those living with others, people who lived alone were far more likely to have low (or negative) incomes. Consistent with previous research,[27] people living by themselves were also significantly more likely than those living with others to have functional disabilities. The group who lived alone was also substantially older than those who cohabited. Not surprisingly, people living alone also included a disproportionately large number of individuals who had never married, were widowed, or were separated or divorced. These patterns likely reflect demographic, psychological, social, and economic factors involved in selection into different living arrangements over the adult lifespan.

Selection and direct causal mechanisms may contribute to the increased suicide risks of adults who live alone. Selection operates through factors that are causally related to living alone and suicide risk. As an example, suicide risk is elevated in the aftermath of divorce and separation,[28] and these transitions also typically result in changes in living arrangements. Because living alone was associated with modest increased risk of suicide among separated or divorced adults in the present report, factors other than living alone such as stress related to separation or divorce[29] or the association of common psychiatric disorders with separation and divorce[30] might also contribute to the elevated risk of suicide among separated or divorced adults.[28,31] The role of selection versus direct mechanisms related to loneliness and social isolation in suicide risk remains unknown. However, the high fraction of adults who live alone that are divorced or separated (34.9%) likely contributes to the high crude rate of suicide among people who live alone.

While beyond the scope of the current analysis, the experience of living alone may also increase suicide risk.

TABLE 2— Suicide Risk of Adults Who Live Alone or With Others Stratified by Sociodemographic Characteristics: Mortality Disparities in American Communities, United States, 2008–2019

Characteristic	Suicide Rate per 100 000 Person-Years: Adults Living Alone (95% CI)	Suicide Rate per 100 000 Person-Years: Adults Living With Others (95% CI)	AHR[a] of Suicide for Living Alone (95% CI) Reference, Living With Others	Interaction (*P*)
Total	23.0 (21.6, 24.4)	13.2 (12.8, 13.6)	1.75 (1.64, 1.87)	
Age, y				
18–39	18.0 (15.7, 20.6)	12.6 (12.0, 13.2)	1.23 (1.07, 1.41)	Ref
40–64	28.7 (26.5, 31.1)	13.4 (12.8, 14.1)	2.15 (1.96, 2.35)	< .001
≥ 65	17.9 (15.7, 20.4)	14.7 (13.4, 16.0)	1.97 (1.68, 2.31)	.001
Sex				
Male	40.3 (37.6, 43.0)	21.0 (20.3, 21.7)	1.82 (1.69, 1.97)	.002
Female	8.9 (7.8, 10.1)	5.8 (5.5, 6.2)	1.69 (1.45, 1.97)	Ref
Race/ethnicity				
Non-Hispanic White	27.6 (25.9, 29.4)	16.1 (15.6, 16.6)	1.79 (1.67, 1.93)	Ref
Non-Hispanic Black	4.7 (3.2, 6.7)	6.1 (5.3, 7.0)	0.92 (0.63, 1.33)	< .001
Hispanic	14.2 (10.5, 18.7)	6.4 (5.7, 7.1)	2.20 (1.63, 2.96)	.26
Other	18.4 (13.1, 24.9)	11.0 (9.6, 12.5)	1.67 (1.20, 2.32)	.61
Marital status				
Married	16.3 (11.2, 22.9)	12.4 (11.9, 12.9)	1.31 (0.93, 1.84)	.53
Separated/divorced	28.2 (25.8, 30.8)	18.1(16.6, 19.6)	1.28 (1.13, 1.46)	.15
Widowed	12.4 (10.4, 14.8)	6.9 (5.2, 9.0)	1.63 (1.15, 2.31)	.36
Never married	24.6 (22.4, 27.0)	13.8 (13.0, 14.7)	1.35 (1.20, 1.51)	Ref
Employment				
Employed	20.8 (19.2, 22.6)	11.4 (11.0, 11.9)	1.82 (1.67, 1.99)	Ref
Not employed, < 65 y	38.2 (33.8, 43.0)	17.9 (16.9, 18.9)	1.74 (1.52, 1.98)	.61
Not employed, ≥ 65 y	18.6 (16.2, 21.4)	15.3 (13.9, 16.8)	1.93 (1.63, 2.29)	.47
Education				
< high school	22.1 (18.5, 26.1)	14.0 (12.9, 15.1)	1.77 (1.46, 2.16)	< .001
High school/GED	22.4 (19.9, 25.2)	15.6 (14.8, 16.4)	1.48 (1.30, 1.69)	< .001
Some college/associate degree	24.8 (22.3, 27.6)	13.4 (12.6, 14.1)	1.82 (1.61, 2.04)	.006
≥ bachelor's degree or higher	22.0 (19.6, 24.6)	10.0 (9.4, 10.8)	2.25 (1.97, 2.56)	Ref
Income, $				
0 to 40 000 (loss)	23.1 (21.4, 24.9)	15.4 (14.5, 16.3)	1.38 (1.26, 1.52)	.014
40 001 to 75 000	23.0 (20.2, 26.0)	13.4 (12.7, 14.2)	1.55 (1.35, 1.77)	.07
75 001 to 125 000	19.8 (15.6, 24.6)	12.2 (11.4, 12.9)	1.45 (1.15, 1.83)	.05
> 125 000	28.1 (20.5, 37.6)	11.3 (10.4, 12.2)	2.22 (1.64, 3.00)	Ref
Residence				
Urban	30.0 (26.4, 34.0)	16.3 (15.4, 17.2)	1.73 (1.60, 1.87)	.59
Rural	21.4 (20.0, 23.0)	12.2 (11.8, 12.7)	1.89 (1.65, 2.17)	Ref
Housing finance				
Renter	22.4 (20.4, 24.4)	11.6 (10.9, 12.3)	1.56 (1.40, 1.74)	.010
Owner	23.5 (21.6, 25.5)	13.8 (13.4, 14.3)	1.83 (1.67, 2.00)	Ref
Residential stability, y				
< 5	23.3 (21.3, 25.4)	12.4 (11.8, 13.0)	1.68 (1.52, 1.86)	Ref
5–10	23.1 (20.2, 26.4)	13.6 (12.8, 14.5)	1.76 (1.52, 2.05)	.73
> 10	22.4 (20.2, 24.9)	13.8 (13.2, 14.6)	1.92 (1.71, 2.17)	.41

Continued

TABLE 2— Continued

Characteristic	Suicide Rate per 100 000 Person-Years: Adults Living Alone (95% CI)	Suicide Rate per 100 000 Person-Years: Adults Living With Others (95% CI)	AHR[a] of Suicide for Living Alone (95% CI) Reference, Living With Others	Interaction (*P*)
Functional disability				
Present	30.4 (27.1, 34.1)	25.5 (23.8, 27.2)	1.49 (1.31, 1.71)	.001
Absent	21.0 (19.6, 22.5)	11.7 (11.3, 12.1)	1.77 (1.64, 1.91)	Ref

Notes. AHR = adjusted hazard ratio; CI = confidence interval; GED = general educational development. Limited to adults aged ≥ 18 years; excludes respondents living in group quarters. Respondents followed through 2019.

[a]Adjusted for age, sex, and race/ethnicity. Disclosure Review Board approval number CBDRB-FY22-CES004-040.

In epidemiological research, living alone has been consistently related to a substantially elevated risk of loneliness,[32] and loneliness has been related to suicidal behavior.[33] Without a measure of loneliness in the present study, however, we were unable to assess the extent to which loneliness, social isolation, other psychological factors, less opportunity for rescue from a suicide attempt, or other factors related to living alone mediate the observed association of living alone with suicide risk.

Although the current study is not intended to evaluate causal connections between living alone and suicide risk, the findings are consistent with a long tradition of sociological research on suicide that has emphasized social disengagement and loss of regulation, related to declining oversight and guidance from social ties.[34] These concepts have their historical roots in Durkheim's insights more than a century ago on the stability of well-integrated groups with cohesive and durable social ties.[35]

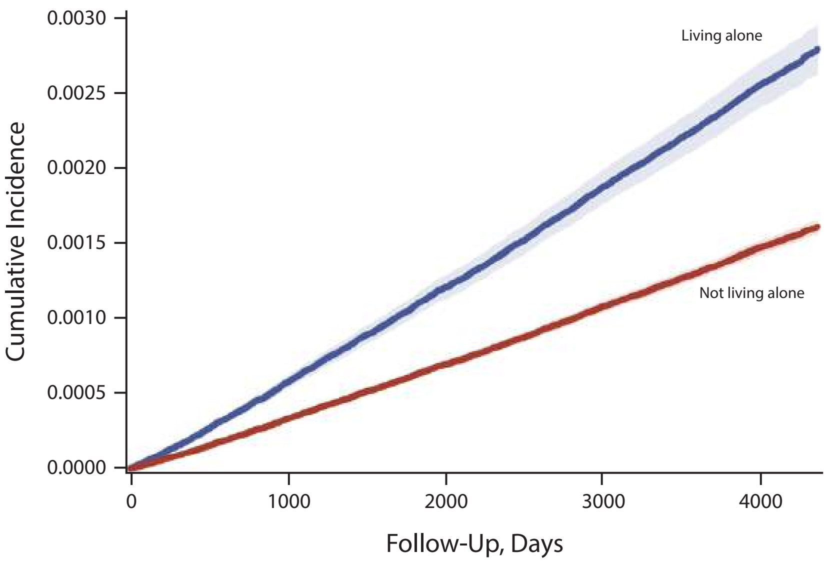

FIGURE 1— Cumulative Suicide Risk of Adults Who Live Alone or With Others: Mortality Disparities in American Communities, United States, 2008–2019

Notes. Analysis was limited to adults aged ≥ 18 years. Disclosure Review Board approval number CBDRB-FY22-CES004-043.

In the current study, living alone was especially strongly related to suicide by poisoning. When poisoning events occur among people who live alone, there may be fewer opportunities for another individual to intercede with a potentially life-saving intervention such as activating the emergency medical services response system.

The current findings suggest that, as a marker of suicide risk, living alone operates differentially across age, sex, ethnic/racial, and educational groups in the United States and underscores opportunities for future research to probe the basis of these variations. For example, the reasons that living alone was not a risk marker for suicide for non-Hispanic Black adults, a group with comparatively low but increasing suicide risk, offers opportunities for research on culturally mediated protective mechanisms. It is possible that strong familial connections among non-Hispanic Black individuals helped to buffer the connection between living alone and suicide risk in this group.[36,37]

Limitations

This analysis had several limitations. First, living arrangements and the other baseline respondent characteristics, especially employment and income, may have changed during follow-up in ways that altered the overall association between living alone and suicide risk and affected its moderation by the sociodemographic characteristics. Although less is known about the stability of living arrangements among younger adults, approximately 81% to 88% of surviving older adults who lived alone at baseline in 2 cohort studies were reported to continue to live alone at 5-year follow-up.[38,39] In the ACS cohort, the 1-year and 11-year follow-up analyses of living alone and suicide risk yielded similar results.

Second, death certificate data may not accurately capture suicide, although suicide in death certificates has been found to have a sensitivity of 90% with information from hospital, autopsy, law enforcement, and medical examiner records as the criterion standard.[40]

Third, because the ACS does not measure important suicide risk factors such as mental health and substance use disorders,[18] previous suicide attempts,[41] or stressful life events[42] that may also be related to living alone,[17,19] the associations between living arrangements and suicide risk should be interpreted as noncausal.

Fourth, the cohort was either not sufficiently large or did not include measures of several other groups with increased rates of suicide including survivors of critical illnesses[43] or individuals who identify as Native Americans[44] or as lesbian, gay, bisexual, transgender, or queer or questioning.[45] Finally, the group who lived with others includes a diverse set of living arrangements that may vary in their associations with suicide risk.[46]

Public Health Implications

The current findings have implications for clinical practice and future epidemiological research. In contrast to loneliness, which is difficult for primary care clinicians to identify in their patients,[47] living alone is a readily discernible personal characteristic. In addition to traditional suicide risk factors, such as depression, substance use, and previous suicidal behavior, clinical consideration might also be given to living circumstances as a risk marker to consider in the context of known suicide risk factors.

The findings might also help inform future research aimed at understanding why the increase in suicide risk among people who live alone varies across sociodemographic characteristics. In this regard, longitudinal designs, which permit probing how transitions in housing arrangements covary with known risk factors for suicide, such as social isolation or depressed mood, might help to elucidate causal mechanisms that contribute to sociodemographic variation in the strength of associations between living alone and death by suicide. AJPH

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CONFLICTS OF INTEREST

The authors report no conflicts of interest.

HUMAN PARTICIPANT PROTECTION

The results were reviewed and approved for release by the US Census Bureau's Disclosure Review Board (DRB) to prevent disclosure of confidential information: DRB releases CBDRB-FY22-CES004-040, CBDRB-FY22-CES004-041, and CBDRB-FY22-CES004-043.

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Trend in Loaded Handgun Carrying Among Adult Handgun Owners in the United States, 2015–2019

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See also Bonne, p. 1705.

Objectives. To determine the frequency of loaded handgun carrying among US adult handgun owners overall and by state concealed carry law status.

Methods. Using a nationally representative survey of US firearm-owning adults in 2019, we asked handgun owners (n = 2389) about their past-month handgun carrying behavior.

Results. A total of 30.3% (95% confidence interval [CI] = 28.0%, 32.6%) of handgun owners carried handguns monthly, of whom 38.1% (95% CI = 33.6%, 42.7%) did so daily. In permitless carry states, 29.7% (95% CI = 25.9%, 33.9%) of handgun owners carried handguns in the past month, compared with 33.1% (95% CI = 29.9%, 36.3%) in shall issue states and 19.7% (95% CI = 14.9%, 25.5%) in may issue states. Of handgun owners without a permit, 7.5% (95% CI = 4.1%, 13.3%) of those in may issue states and 11.5% (95% CI = 8.5%, 15.4%) of those in shall issues states carried handguns in the past month.

Conclusions. In 2019, about 16 million US adult handgun owners carried handguns in the past month (up from 9 million in 2015), and approximately 6 million did so daily (twice the 3 million who did so in 2015). Proportionally fewer handgun owners carried handguns in states where issuing authorities had substantial discretion in granting permits. (*Am J Public Health*. 2022;112(12):1783–1790. https://doi.org/10.2105/AJPH.2022.307094)

AJPH December 2022, Vol 112, No. 12

Little is known about the frequency and features of firearm carrying among adult handgun owners in the United States. In fact, over the past 30 years, only a few peer-reviewed national surveys, conducted in 1994, 1995, 1996, and 2015, have provided even the most basic information about firearm carrying frequency.[1–4] Since the first of these surveys, reasons offered by firearm owners for why they own firearms have shifted from hunting and sports shooting toward personal protection. In 1994, for example, 46% of firearm owners reported owning firearms for protection[2]; by 2015, that number had reached 65%,[5] and, by 2019, it had reached 73%.[6] As personal protection became the predominant motivation for owning firearms, handgun ownership increased disproportionately from 64% in 1994 to 83% in 2021.[2,7]

These trends have been accompanied by a loosening of state laws governing who can carry handguns in public places. State laws regulating concealed handgun carrying are typically divided into the following types: (1) permitless: no permit is required; (2) shall issue: the issuing authority is required to grant a permit to anyone who meets certain minimal statutory requirements with no or limited discretion; (3) may issue: the issuing authority has substantial discretion to approve or deny a concealed carry permit to an applicant.[8] In 1990, only 1 state allowed permitless handgun carry; at the time of this writing, that number had risen to 21.[8]

To our knowledge, the only contemporary national estimates of handgun carrying among US adults come from the National Firearms Survey in 2015 (NFS-2015). NFS-2015 found that 23.5% of adult handgun owners (9 million adults) had carried a loaded handgun on their person in the month before the survey; of those, 34.5% (3 million) had done so every day.[4] Of handgun owners who carried, 4 in 5 carried primarily for protection, 4 in 5 had a concealed carry permit, 2 in 3 always carried concealed, and 1 in 10 always carried openly.[4]

The prevalence of handgun carrying was similar in states with permitless carry laws and states with shall issue carry laws. By contrast, the prevalence of carrying was notably lower in states with may issue carry laws.[4]

In the current study (NFS-2019), we used nationally representative survey data collected from July 30, 2019, to August 11, 2019, to update information pertaining to the proportion of handgun owners who carried a handgun over the previous month (and, of those, the fraction who carried daily), the characteristics of those who carried, and the prevalence of handgun carrying by handgun owners in states that did versus did not require a permit for concealed carrying at the time of the survey.

METHODS

Data for this cross-sectional study came from the Web-based NFS-2019. We designed the survey to assess firearm-related beliefs and behaviors, including handgun carrying, in a nationally representative sample of US adults living in firearm-owning households. The survey was conducted by the research firm Ipsos from July 30, 2019, to August 11, 2019. Consistent with NFS-2015, respondents were drawn from Ipsos's Knowledge Panel, an online sampling frame comprising approximately 55 000 US adults selected using address-based sampling methods on an ongoing basis with an equal probability of selection.

Panel members' report of whether they live in a home with firearms is collected on enrollment in Knowledge Panel and updated approximately annually, allowing us to restrict invitations for participation to adults (other than those on active duty in the US military) who reported that they lived in a home with firearms. E-mail invitations to participate in the survey contained a link that sent them to the survey questionnaire. No description of the survey content accompanied the invitation. Reminder e-mails were sent to nonresponders on days 3, 6, 9, and 12. Ipsos has a modest point-based incentive program through which participants accrue points to redeem for rewards. Of the 6721 panel members invited to complete the survey, 4379 started and 4030 completed the survey (response proportion: 65.2%; participation proportion: 92.0%). Participants were not involved in the design, conduct, reporting, or dissemination plans of our research. More details about the survey can be found elsewhere.[9]

Measures

Firearm ownership status was determined by responses to the question: "Do you personally own a gun?" Only respondents who responded affirmatively were then asked questions about the type of firearms owned. The current study was limited to respondents within firearm-owning households who reported that they personally owned a handgun ($n = 2389$) regardless of whether they also owned a long gun. The survey focused exclusively on loaded handgun carrying on the person, and not in a vehicle. Respondents were asked: "In the past 30 days, have you carried a loaded handgun on your person?" Those who answered affirmatively were then asked about the number of days that they had carried (range = 0–30) and the primary reason for carrying.

Additional survey domains included respondents' sociodemographic characteristics, presence of children in the home, veteran status, type of firearm owned (handguns only vs handguns and long guns), and holding a concealed carry permit. Selected survey questions related to this analysis are provided in the Appendix (available as a supplement to the online version of this article at https://ajph.org). State handgun carry laws were identified using the state law database at Giffords Law Center to Prevent Gun Violence.[8] State laws were coded by whether they required a permit for concealed handgun carrying in public in July to August 2019.

Some survey respondents refused to answer some questions about handgun ownership and handgun carrying behaviors. Of the 4030 total respondents, 1 refused to answer if they had carried a loaded handgun in the past 30 days, 3 refused to answer how many days they had carried a loaded handgun, 28 refused to answer if they had a concealed carry permit, and 15 refused to answer questions about the types of guns they owned (handguns, long guns, etc.). In addition, 81 respondents said that they did not know if they had a concealed carry permit. These refused-to-answer or "do not know" responses were recoded as missing.

A total of 20 respondents who indicated that they had carried a handgun in the past 30 days responded "0 days" when asked for the number of days in which they had carried their handgun in the past 30 days. We recoded these individuals as "did not carry in the past 30 days" in our analyses. There was no missingness in information on the reason for carrying a handgun. We generated variables about carrying laws by state; because there was no missingness in the state data for respondents, there was no missingness in carrying law variables either. There was no missingness in any demographic data from the survey respondents. Overall, because of the low frequency of missing values in this analysis, we conducted no imputation.

Statistical Analysis

Analyses used individual-level survey weights provided by Ipsos. These weights account for survey nonresponse and under- or overcoverage imposed by the study-specific sample design. Weights also adjusted for benchmark demographic distributions from the US Census Current Population Survey or the American Community Survey and population characteristics that are not available in either of those surveys (e.g., firearm ownership) based on Knowledge Panel profile data for gender, age, race, ethnicity, census region, metropolitan statistical area status, and education.

For this analysis, we calculated weighted percentages and their corresponding 95% confidence intervals (CIs) for each measure. Sociodemographic characteristics of handgun owners and type of firearms owned were described by past-30-day handgun carrying status (i.e., did carry vs did not carry). We conducted all analyses in Stata version 14 (StataCorp LP, College Station, TX) using the "svy" suite of commands.

RESULTS

Of all handgun owners, 30.3% (95% CI = 28.0%, 32.6%) reported having carried a handgun in the past 30 days. Among those, 38.1% (95% CI = 33.6%, 42.7%) reported doing so every day (Figure 1). Among handgun owners who reported carrying at least once in the past 30 days, the mean number of carrying days was 18.1 (95% CI = 17.1, 19.2).

Compared with handgun owners who did not carry a handgun, a significantly greater proportion of those who carried were younger, male, lived in the South (South Atlantic, East–South Central, West–South Central), and owned both handguns and long guns (Table 1). The distributions of race, educational attainment, annual household income, urbanicity of the community of residence, presence of children in the

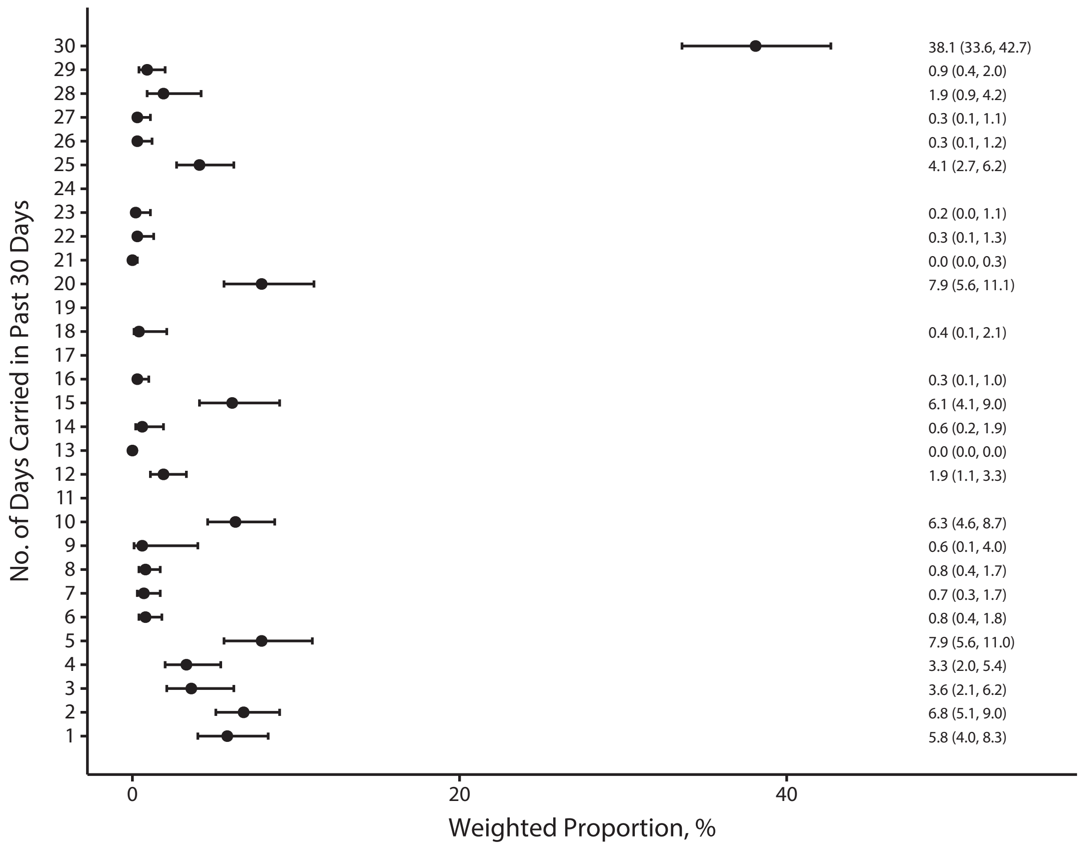


FIGURE 1— Number of Carrying Days Among Handgun Owners Who Carried a Handgun in the Past 30 Days: National Firearms Survey, United States, 2019

Note. Horizontal bars in the figure are 95% confidence intervals.

TABLE 1— Characteristics of Handgun Owners by Past-30-Day Carrying Status: National Firearms Survey, United States, 2019

Characteristics	No.	Did Not Carry (n = 1736), % (95% CI)	Carried (n = 653), % (95% CI)	Total (n = 2389), % (95% CI)	P
All respondents	2389	69.7 (67.4, 72.0)	30.3 (28.0, 32.6)	100	
Age, y					< .001
18–29	121	9.4 (7.4, 11.9)	16.9 (13.0, 21.8)	11.7 (9.7, 13.9)	
30–44	433	22.3 (19.8, 24.9)	24.4 (20.7, 28.6)	22.9 (20.8, 25.1)	
45–59	709	30.5 (28.0, 33.1)	32.8 (28.6, 37.3)	31.2 (29.0, 33.5)	
≥ 60	1126	37.8 (35.3, 40.4)	25.9 (22.5, 29.5)	34.3 (32.2, 36.4)	
Gender					< .001
Male	1651	63.4 (60.6, 66.1)	79.3 (75.4, 82.8)	68.2 (65.9, 70.4)	
Female	738	36.6 (33.9, 39.4)	20.7 (17.2, 24.6)	31.8 (29.6, 34.1)	
Race/ethnicity					.31
Non-Hispanic White	1968	77.6 (74.8, 80.1)	72.8 (68.0, 77.1)	76.1 (73.7, 78.4)	
Non-Hispanic Black	159	8.6 (6.9, 10.6)	10.6 (7.8, 14.2)	9.1 (7.7, 10.9)	
Non-Hispanic other	39	3.0 (2.0, 4.5)	4.6 (2.6, 8.0)	3.5 (2.5, 4.9)	
Hispanic	148	9.1 (7.3, 11.2)	10.5 (7.6, 14.4)	9.6 (8.0, 11.4)	
Non-Hispanic ≥ 2 races	75	1.8 (1.2, 2.5)	1.5 (0.9, 2.5)	1.7 (1.3, 2.3)	
Education					.61
Less than high school	69	6.0 (4.4, 8.0)	5.3 (3.2, 8.6)	5.7 (4.4, 7.4)	
High school	480	28.4 (25.8, 31.3)	29.2 (24.8, 34.0)	28.7 (26.4, 31.1)	
Some college	826	32.5 (29.9, 35.2)	35.5 (31.3, 39.9)	33.5 (31.2, 35.7)	
Bachelor's degree or higher	1014	33.1 (30.6, 35.6)	30.1 (26.3, 34.1)	32.1 (30.1, 34.3)	
Annual household income, $					.6
< 25 000	183	7.6 (6.3, 9.3)	8.3 (6.0, 11.5)	7.9 (6.6, 9.3)	
25 000–74 999	826	32.8 (30.2, 33.5)	35.6 (31.2, 40.3)	33.7 (31.5, 36.0)	
75 000–124 999	752	30.0 (27.5, 32.6)	29.4 (25.5, 33.7)	29.8 (27.6, 32.0)	
≥ 125 000	628	29.5 (27.0, 32.2)	26.7 (22.7, 31.0)	28.7 (26.5, 31.0)	
Community of residence					.11
Metro area	1929	81.5 (79.1, 83.6)	77.9 (73.7, 81.5)	80.4 (78.4, 82.3)	
Non–metro area	460	18.5 (16.4, 20.9)	22.1 (18.5, 26.3)	19.6 (17.7, 21.6)	
Region based on residence[a]					.024
New England	62	2.2 (1.6, 3.1)	3.8 (2.2, 6.3)	2.7 (2.0, 3.6)	
Mid-Atlantic	194	7.9 (6.5, 9.5)	7.1 (5.2, 9.5)	7.6 (6.5, 8.9)	
East–North Central	344	13.3 (11.6, 15.3)	11.8 (9.2, 15.0)	12.9 (11.4, 14.5)	
West–North Central	210	8.1 (6.8, 9.7)	6.5 (4.6, 9.1)	7.7 (6.6, 9.0)	
South Atlantic	491	19.8 (17.6, 22.1)	25.0 (21.3, 29.1)	21.4 (19.5, 23.4)	
East–South Central	190	8.0 (6.6, 9.7)	9.8 (7.3, 13.1)	8.5 (7.2, 10.0)	
West–South Central	330	15.9 (13.7, 18.2)	16.4 (12.9, 20.6)	16.0 (14.1, 18.0)	
Mountain	266	11.3 (9.6, 13.3)	11.3 (8.6, 14.7)	11.3 (9.8, 13.0)	
Pacific	302	13.5 (11.7, 15.5)	8.4 (6.4, 11.0)	12.0 (10.6, 13.6)	
Children (< 18 y) in household					.16
None	1828	69.6 (66.7, 72.3)	65.8 (61.1, 70.2)	68.5 (66.0, 70.8)	
≥ 1	561	30.4 (27.7, 33.3)	34.2 (29.8, 38.9)	31.5 (29.2, 34.0)	
Veteran					.3
Yes	573	19.7 (17.7, 21.9)	21.9 (18.4, 25.7)	20.4 (18.6, 22.3)	
No	1816	80.3 (78.1, 82.3)	78.1 (74.3, 81.6)	79.6 (77.7, 81.4)	

Continued

TABLE 1— Continued

Characteristics	No.	Did Not Carry (n = 1736), % (95% CI)	Carried (n = 653), % (95% CI)	Total (n = 2389), % (95% CI)	P
Firearm type owned					< .001
Handgun only	900	44.3 (41.5, 47.2)	24.8 (20.9, 29.0)	38.3 (36.0, 40.7)	
Handgun and long gun	1489	55.7 (52.8, 58.5)	75.2 (71.0, 79.1)	61.7 (59.3, 64.0)	

Note. CI = confidence interval. Column percentages are weighted sample proportions.

[a]New England comprises CT, ME, MA, NH, RI, and VT. Mid-Atlantic comprises NJ, NY, and PA. East–North Central comprises IL, IN, MI, OH, and WI. West–North Central comprises IA, KS, MN, MO, NE, ND, and SD. South Atlantic comprises DE, FL, GA, MD, NC, SC, VA, and WV. East–South Central comprises AL, KY, MS, and TN. West–South Central comprises AR, LA, OK, and TX. Mountain comprises AZ, CO, ID, MT, NV, NM, UT, and WY. Pacific comprises AK, CA, HI, OR, and WA. *P* values are calculated based on a design-based F-test that takes the sampling design into account.

household, and veteran status were not notably different between handgun owners who carried and those who did not (Table 1). Prevalence estimates of handgun carrying by specific handgun owner characteristics are available in Figure A (available as a supplement to the online version of this article at https://ajph.org). Most handgun owners who carried a handgun did so primarily for personal protection against people (71.8%; 95% CI = 67.4%, 75.8%; Figure 2). Findings on the frequency of and reasons for carrying stratified by gender are found in Figures B through E (available as supplements to the online version of this article at https://ajph.org).

A smaller proportion of handgun owners residing in may issue states (19.7%; 95% CI = 14.9%, 25.5%) carried a handgun than did those in permitless carry (29.7%; 95% CI = 25.9%, 33.9%) and shall issue (33.1%; 95% CI = 29.9%, 36.3%) states (Figure 3). Approximately 11.0% (95% CI = 8.7%, 13.8%) of handgun owners who lived in a state that required a permit to carry but did not themselves have a permit reported that they had carried a handgun in the past month. Specifically, of handgun owners who resided in a may issue state but did not have a permit, 7.5% (95% CI = 4.1%, 13.3%) carried a handgun; of handgun owners who resided in a shall issue state but did not have a permit, 11.5% (95% CI = 8.5%, 15.4%) carried a handgun (Figure F, available as a supplement to the online version of this article at https://ajph.org).

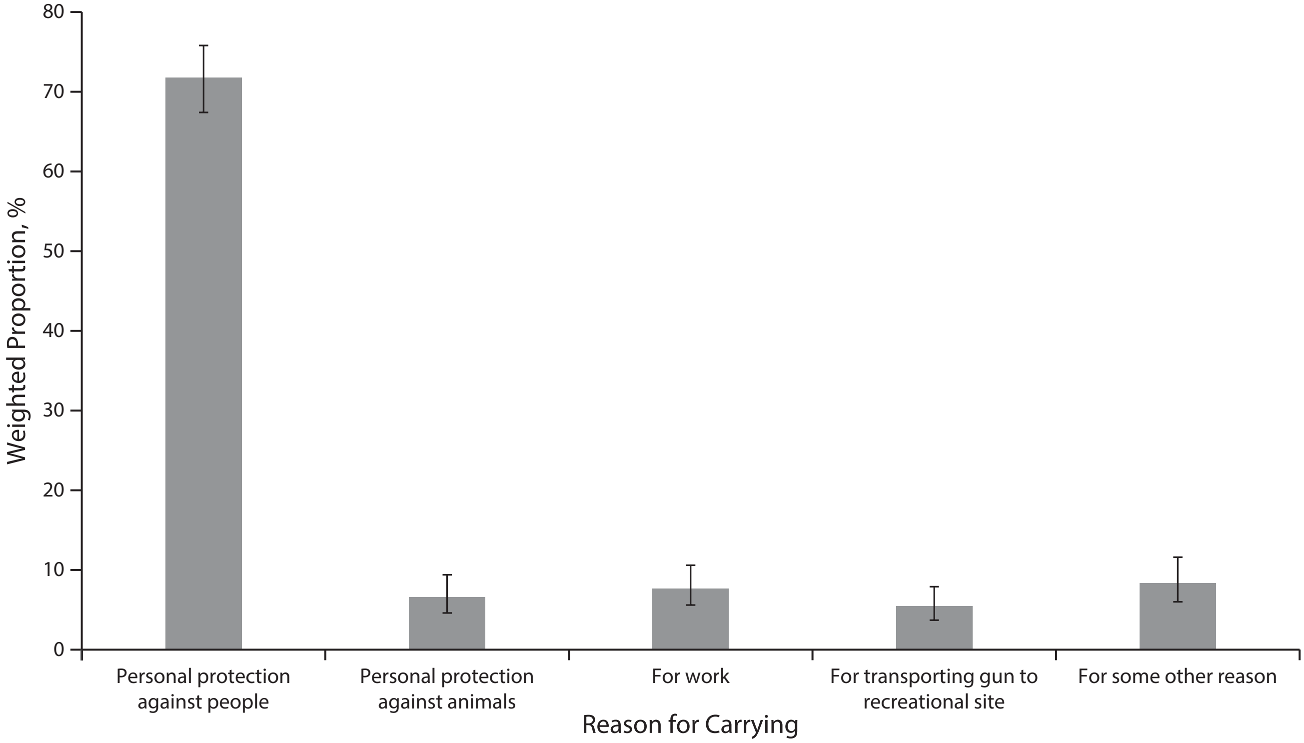
DISCUSSION

In this nationally representative study conducted in 2019 (NFS-2019), we found that about 3 in 10 handgun owners carried a loaded handgun on their person in the past 30 days; among those, about 4 in 10 did so every day. Extrapolating to the estimated 53 million US adults who owned handguns in 2019, we estimate that about 16 million US adults carried a handgun in the past 30 days (up from 9 million in 2015), and that almost 6 million did so every day (twice the approximately 3 million who did so in 2015).[4,7]

In NFS-2019, about 7 in 10 handgun owners who carried handguns cited protection against people as the main reason for carrying. This proportion was greater (about 8 in 10) in NFS-2015. This difference could indicate an actual decline in the proportion of handgun owners who carried for personal protection against people from 2015 to 2019 or, alternatively, could be attributable to differences in the wording of the questions asked in 2015 and 2019. In 2015, the options included "For protection against strangers" and "For protection against people I know" while in 2019 the option was "For personal protection against people" (other options in both surveys included for protection against animals, hunting, sporting, and other reasons). Regardless, results from the current survey continue to demonstrate that a large majority of handgun owners who carry do so for self-defense.

We found no notable differences between the proportion of handgun owners residing in permitless carry states who carried handguns versus those residing in shall issue states who did so. Consistent with findings from NFS-2015, however, we found that proportionally fewer handgun owners residing in may issue states than those residing in permitless carry states and shall issue states carried handguns in 2019. In 2015, we found that 21.1% and 9.1% of handgun owners residing in permitless states and may issue states at that time had carried handguns, respectively. In 2019, those numbers were 33.1% and 19.7%, respectively.

In addition, in 2015, only 1.2% of handgun owners without a permit residing in may issue states had carried handguns; that number rose to 7.5%



Personal protection against people	Personal protection against animals	For work	For transporting gun to recreational site	For some other reason
71.8% (95% CI = 67.4%, 75.8%)	6.6% (95% CI = 4.6%, 9.4%)	7.7% (95% CI = 5.6%, 10.6%)	5.5% (95% CI = 3.7%, 7.9%)	8.4% (95% CI = 6.0%, 11.6%)

FIGURE 2— Primary Reason for Handgun Carrying Among Handgun Owners Who Carried a Handgun in the Past 30 Days: National Firearms Survey, United States, 2019

Note. Vertical bars in the figure are 95% confidence intervals (CIs).

in 2019. The NFS-2015 question specifically asked about concealed carrying whereas the NFS-2019 question asked about carrying. Nonetheless, if in 2019, as in 2015, only 10% of handgun owners always carried handguns openly (and thus would not necessarily be in violation of a permit law),[4] our findings still suggest a substantial increase in the number of handgun owners who carried handguns without a permit when they were legally required to have one.

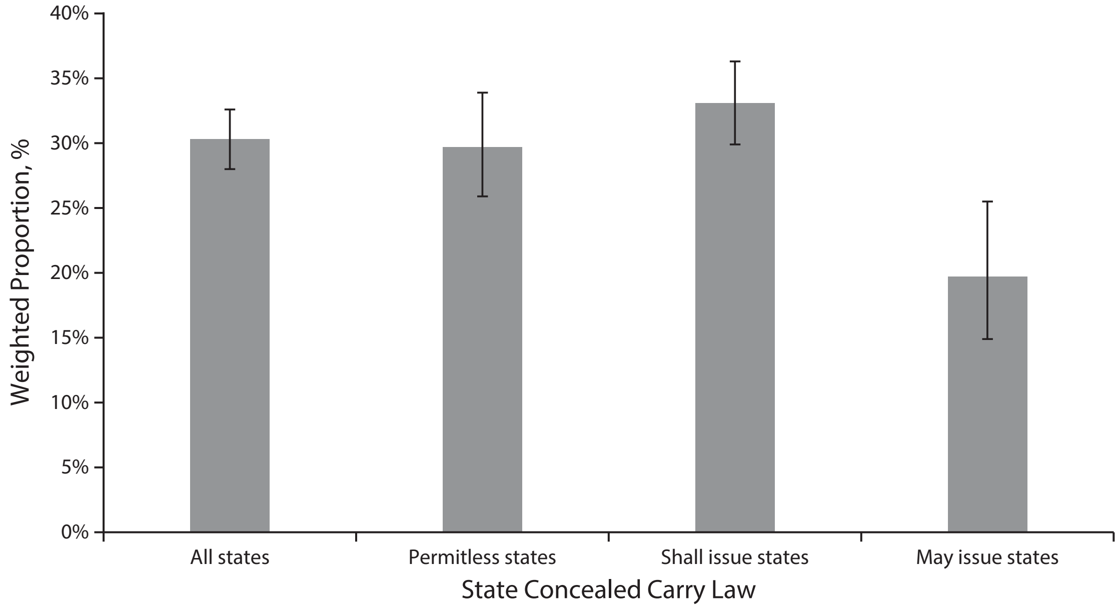
Limitations

Our study was subject to limitations. First, we did not ask survey respondents in which state they had carried their handgun; they may have carried their handgun in a state different from the one in which they resided at the time of the survey resulting in some degree of misclassification in our findings pertaining to the prevalence of handgun carrying in relation to state laws.

Second, NFS-2019 did not ask respondents whether they carried a handgun concealed or openly. However, it is likely a safe assumption that that the overwhelming majority of those who carried handguns did so concealed, at least on some days. In NFS-2015, for example, only about 10% of handgun owners who carried said they always carried openly. If that same fraction pertained in 2019, it would revise our estimate of the number of past-month carriers to 14.6 million, which still represents a substantial increase from 2015.[4]

Third, as in all self-report surveys, recall and reporting bias may have affected our results. To minimize recall error, questions on handgun carrying referred to the 30-day period before the survey, reducing concerns about recall bias. Although reporting bias (e.g., social desirability bias) may still have affected our results, online panel surveys such as ours tend to be less biased than alternatives, such as telephone surveys, in this specific aspect.[10]

Fourth, panel members who chose not to participate in our survey may have been different from those who



All states	Permitless states	Shall issue states	May issue states
30.3% (95% CI = 28.0%, 32.6%)	29.7% (95% CI = 25.9%, 33.9%)	33.1% (95% CI = 29.9%, 36.3%)	19.7% (95% CI = 14.9%, 25.5%)

FIGURE 3— Prevalence of Past-30-Day Handgun Carrying Among Handgun Owners by State Concealed Carry Laws: National Firearms Survey, United States, 2019

Note. Vertical bars in the figure are 95% confidence intervals (CIs).

chose to participate regarding their frequency and features of handgun carrying. However, an advantage of online panels is high completion proportions for individuals who begin the survey.[11] In our study, the completion proportion was 92.0%; only 15 respondents refused to answer handgun ownership questions, and only 4 respondents refused to answer handgun carrying behavior questions. Our survey response proportion of 65.2% is also substantially greater than the range of percentages observed in typical nonprobability, opt-in, online surveys (2%–16%).[11]

Public Health Implications

On November 3, 2021, the US Supreme Court heard its first case explicitly related to handgun carrying (*New York State Rifle & Pistol Association v. Bruen*).[12] The case tested whether the New York law requiring lawful firearm owners to provide a proper cause to obtain a permit to carry is too restrictive. On June 23, 2022, the Supreme Court ruled that New York's proper-cause requirement violates the Fourteenth Amendment's guarantee of equal protection under the law by preventing law-abiding citizens with ordinary self-defense needs from exercising their Second Amendment right to keep and bear arms in public for self-defense.[13] This ruling could further catalyze the loosening of firearm-carrying regulations in different parts of the country at a time when, as our study indicates, trends in handgun carrying already point to more US adults carrying loaded handguns in public places, including without a permit when a permit is required. The effect of this loosening on firearm ownership and carrying as well as public safety and public health should be an important subject of research in the future. AJPH

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CONTRIBUTORS

A. Rowhani-Rahbar, D. Azrael, and M. Miller contributed to the conceptualization and design of the study. All authors contributed to the acquisition, analysis, or interpretation of data. A. Rowhani-Rahbar drafted the article. All authors contributed to the critical revision of the article for important intellectual content. A. Gallagher conducted the statistical analyses. D. Azrael and M. Miller obtained funding. A. Rowhani-Rahbar, D. Azrael, and M. Miller provided administrative, technical, and material support, and supervision.

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CONFLICTS OF INTEREST

The authors have no conflicts of interest to declare.

HUMAN PARTICIPANT PROTECTION

This study was approved by the institutional review board at the Harvard T. H. Chan School of Public Health.

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School District Prevention Policies and Risk of COVID-19 Among In-Person K–12 Educators, Wisconsin, 2021

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See also Teasdale and Fleary, p. 1696.

Objectives. To assess the rate of COVID-19 among in-person K–12 educators and the rate's association with various COVID-19 prevention policies in school districts.

Methods. We linked actively working, in-person K–12 educators in Wisconsin to COVID-19 cases with onset from September 2 to November 24, 2021. A mixed-effects Cox proportional hazards model, adjusted for pertinent person- and community-level confounders, compared the hazard rate of COVID-19 among educators working in districts with and without specific COVID-19 prevention policies.

Results. In-person educators working in school districts that required masking for students and staff experienced 19% lower hazards of COVID-19 than did those in districts without any masking policy (hazard ratio = 0.81; 95% confidence interval = 0.72, 0.92). Reduced COVID-19 hazards were consistent and remained statistically significant when educators were stratified by elementary, middle, and high school environments.

Conclusions. In Wisconsin's K–12 school districts, during the fall 2021 academic semester, a policy that required both students and staff to mask was associated with significantly reduced risk of COVID-19 among in-person educators across all grade levels. (*Am J Public Health*. 2022;112(12):1791–1799. https://doi.org/10.2105/AJPH.2022.307095)

AJPH December 2022, Vol 112, No. 12

Evidence supports the use of specific prevention efforts to reduce COVID-19 transmission in schools during periods of high community transmission. Policies related to masking,[1–5] physical distancing,[6,7] and quarantine after close contact (resulting from effective contact tracing)[8] have been associated with reduced rates of COVID-19 transmission and outbreaks in school environments. In districts practicing a multifaceted combination of these policies, students and staff experience rates of COVID-19 lower than those of the surrounding communities.[9,10]

For the 2021–2022 academic year, most K–12 students and educators in the United States returned to in-person school environments. In Wisconsin, the vast majority of regular K–12 school districts offered in-person learning for the 2021–2022 school year and were responsible for implementation of their own COVID-19 prevention policies. There was no standardized return-to-school directive from the state regarding implementation of such policies.[11]

The resulting heterogeneity in school district COVID-19 prevention policies throughout Wisconsin allowed us to build on a significant limitation of previous research in this field. Most school-related policy research was conducted during the previous 2020–2021 academic year—a time when almost all schools or districts had some form of prevention policy in place; as a result, it was challenging to directly compare the risk of COVID-19 associated with the presence or absence of certain policies.

In this analysis, our aim was to assess the rate of COVID-19 among in-person K–12 educators and the rate's association with a COVID-19 prevention policy's presence or absence. We chose to compare the rates of COVID-19 among in-person K–12 educators specifically because this is a group that is just as often involved in school-based COVID-19 transmission events as are students[12,13] and is an occupational category with frequently overlooked workplace risk.[14,15] We further stratified educators based on grade level taught to investigate the effect of COVID-19 prevention policies in elementary, middle, and high school settings.

METHODS

We completed our analysis using a cohort study design and a variety of data sources collected prospectively or at a single time point. We used multiple data sources available at both the Wisconsin state and the national levels to aggregate information related to our study sample (Wisconsin's in-person K–12 educators), outcome (COVID-19 cases), and exposure (school district COVID-19 prevention policies).

Educator Data

We created a roster of all licensed, actively working educators in Wisconsin during the 2020–2021 school year from multiple data sources maintained by the Wisconsin Department of Public Instruction. We filtered a data set consisting of all licensed educators in Wisconsin using a data set of educators actively employed during the 2020–2021 school year (the most recent academic year available).[16] We used this merged data set to represent all licensed educators likely to be working during the 2021–2022 school year.

Many categories of educators in Wisconsin can be licensed, including administrators, pupil service staff, and classroom teachers.[16] For educators with multiple categories assigned, we categorized individuals based on their position with the highest full-time equivalent value. We also used these positions to categorize educators by grade level taught (elementary school, middle school or junior high school, and high school). We excluded educators assigned to roles not likely to be working in school settings (Figure 1).

COVID-19 Case Data

We used the Wisconsin Electronic Disease Surveillance System (WEDSS) to collect all confirmed and probable cases of COVID-19 reported from June 1 through November 30, 2021 throughout Wisconsin. We based criteria for confirmed and probable cases on

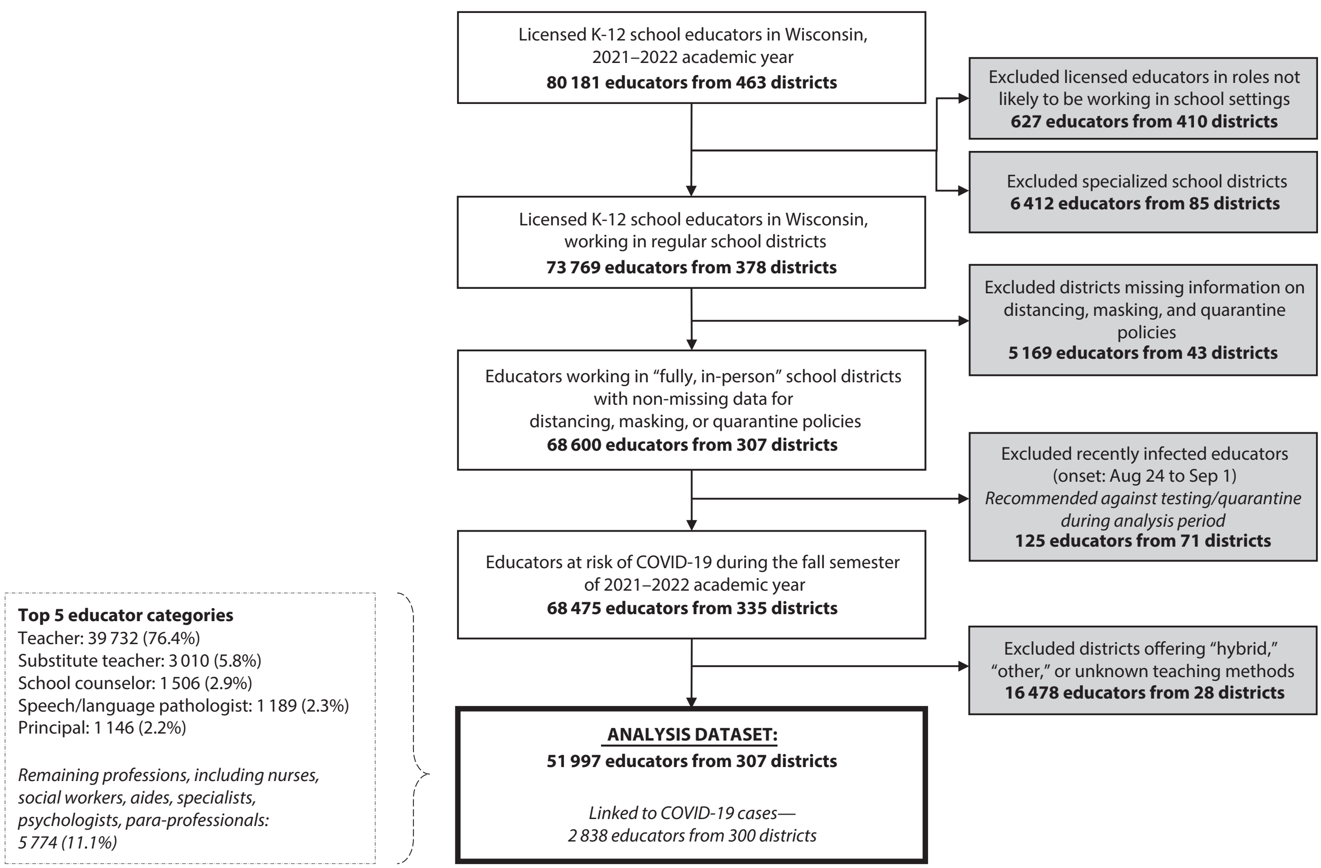


FIGURE 1— Criteria for School District and Educator Inclusion in Study Analysis: Wisconsin, September 2–November 24, 2021

definitions established by the Council for State and Territorial Epidemiologists.[17] We used illness symptom onset date to time stamp cases; if the reported symptom onset date was missing (e.g., for asymptomatic persons), we used the specimen sample collection date or the diagnosis date as a substitute.

September 1 was the first day of the academic year for Wisconsin K–12 schools in 2021. Therefore, the only COVID-19 cases we considered were those that were time stamped from September 2 through November 24, 2021 (or 1 full day into the academic year through the day before Thanksgiving break). We used identifying information from the educator licensure database, including name and date of birth, to link Wisconsin educators to these time-eligible COVID-19 case records in WEDSS. For all cases, we adjusted time at risk during the study period based on the US Centers for Disease Control and Prevention (CDC) recommendation at the time against retesting or quarantine after close contact for persons with infections 90 or fewer days ago.[18] Educators linked to a case of COVID-19 contributed no risk time during their respective 90-day window after infection.

School District Prevention Policy Data

There were 463 school districts in Wisconsin that were linked to our roster of actively working educators in 2021–2022 (Figure 1). Of these, 378 districts were considered "regular school districts" by the National Center for Education Statistics (NCES), which are defined as "locally governed agenc[ies] responsible for providing free public elementary or secondary education."[19] The other 85 districts exist within these regular school districts and generally reflect individual schools or specialized programs (e.g., preparatory academies, schools for deaf or blind students). For the sake of broader generalizability and to avoid issues with small numbers in our results, we excluded these smaller 85 districts.

We obtained COVID-19 prevention policy data for Wisconsin school districts from responses to a national cross-sectional telephone survey.[11] MCH Strategic Data (Sweet Springs, MO) designed the questionnaire in partnership with Esri (Redlands, CA) and the CDC Foundation (Atlanta, GA). This questionnaire was administered to US K–12 public school districts before the start of the 2021–2022 school year. For this analysis, we extracted Wisconsin school district survey responses related to masking, physical distancing, and quarantine policies. The original survey requested specific responses about whether the policy applied to students and educators separately. We operationalized each of these policies as (1) robust—required for both students and educators, (2) partial—required for either students or educators, or (3) absent—required for neither students nor educators (Table A, available as a supplement to the online version of this article at http://www.ajph.org). We excluded districts missing information for all distancing, masking, and quarantine policies.

To adjust for potential ascertainment bias owing to regular COVID-19 testing policies in schools (wherein districts with prevention policies might have also been asking educators to routinely test for COVID-19), we also extracted information on regular staff testing policies.

Person-Level Confounders

We included 3 educator-level variables as potential confounders: age, sex, and COVID-19 vaccination status. We obtained age and sex from the educator licensure information. We collected COVID-19 vaccination information from the Wisconsin Immunization Registry, which the Wisconsin Department of Health Services stores and maintains. We linked educators to COVID-19 vaccination records based on an exact match for first name, last name, and date of birth. We implemented a subsequent linking step using an exact match for date of birth and approximate text matching on both first name and last name. Approximate text matching was based on Jaro–Winkler distance calculations (with distance ≤ 0.25).[20]

Community-Level Confounders

We considered 2 community-level variables to be potential confounders given their association with educator risk outside the school environment and their likely association with COVID-19 policies implemented in school districts. First, we aggregated COVID-19 case data from WEDSS by week for each Wisconsin school district community (i.e., the general population living in school district boundaries), which we used to account for temporal changes in COVID-19 incidence.[21] Second, we accounted for the proportion of the school district community vaccinated against COVID-19 using publicly available Wisconsin Immunization Registry data.[21]

School District–Level Confounders

We incorporated 2 district-specific variables into our analysis as confounders. For one, we calculated a proxy for average classroom size using a student to educator ratio derived from the NCES Common Core data set. Using this same data set, we included the NCES locale classification of school district (city, suburb, town, or rural). Definitions for each locale were based on census-defined groupings and are available on the NCES Web site.[19]

Statistical Analysis

To compare unadjusted differences in school districts with different prevention policies, we used nonparametric statistical tests, including Wilcoxon rank–sum for continuous variables and χ^2 for categorical variables.

To compare hazard rates of COVID-19 among educators working in districts with various prevention policies, we used a mixed-effects Cox proportional hazards model. We adjusted this model for previously described confounders at the individual, community, and school district levels. We included a random effect for school district to account for additional unknown or unobserved confounders at the school district level. We chose to keep all 3 prevention policies in the same multivariate-adjusted regression model to assess their independent contribution to the overall association. We assessed Schoenfeld residuals to confirm that neither the model overall nor the 3 main policy variables violated the proportional hazards assumption.[22] We used spline terms for continuous confounders to allow a nonlinear relationship with the outcome.

We used 4 distinct regression models to account for school districts that were missing district-level data for 1 or 2 COVID-19 prevention policies. Model A included information only from school districts with complete data for all 3 policies. Model B imputed missing policy information using information from nonmissing district-level characteristics, including district population size, proportion of district vaccinated in fall 2021, NCES locale (i.e., urban vs rural), and number of educators and students.[23] As a sensitivity analysis, we created 2 other data sets in which missing policy information was assumed to be either absent (model C) or robust (model D). We conducted all analyses in R version 4.1 (R Foundation for Statistical Computing, Vienna, Austria).[24]

RESULTS

Of the 378 Wisconsin K–12 regular school districts, 43 districts (11.4%) did not submit any response for policies related to physical distancing, mask use, or quarantine (Figure 1). We excluded these districts from our analysis, including the 5169 educators affiliated with them. We also excluded educators who were not considered to be at risk for COVID-19 because a 90-day window following recent infection extended throughout our entire analysis period (n = 125; illness onset dates: August 24–September 1, 2021). Lastly, we excluded all school districts that reported a teaching method for fall 2021 other than “full in-person learning” (n = 28 districts; n = 16 478 affiliated educators). We were left with 51 997 licensed, in-person K–12 educators from 307 school districts in our study sample.

Educators were on average aged 44.0 years; the majority were female (n = 38 702; 74.4%), non-Hispanic White (50 478; 97.1%), and employed by their school district as a teacher (39 732; 76.4%). As of the first day of the 2021–2022 school year (September 1, 2021), 40 526 (77.9%) educators had completed a full primary series of a COVID-19 vaccination. From September 2 through November 24 (the day before the start of Thanksgiving break), 2838 (5.5%) of 51 997 educators were linked to a case of COVID-19. This translated to an unadjusted cumulative incidence of 5458 cases per 100 000 educators.

Responding K–12 public school districts implemented a variety of prevention practices, but policies were nearly always applied to students and staff equally (Table 1; Figure 2). Very few districts implemented a partial policy. Among districts that reported a robust policy practice, physical distancing procedures were the most commonly reported (188/278; 67.6%), followed by quarantine (87/169; 51.5%), and then masking requirements (73/298; 24.5%).

Unadjusted Kaplan–Meier curves indicated that educators working in districts with a robust distancing, masking, or quarantine policy (compared with those working in districts without these policies) experienced a significantly lower hazard of COVID-19 illness from September 2 through November 24, 2021 (Figure A, available as a supplement to the online version of this article at http://www.ajph.org).

Using our imputed multivariate mixed-effects proportional hazards model, we found that, compared with those in districts without masking policies, educators working in districts with robust masking policies were associated with a 19% lower hazard of COVID-19 during September 2 through November 24 (hazard ratio [HR] = 0.81; 95% confidence interval [CI] = 0.72, 0.92). Neither quarantine nor distancing policies were significantly associated with educator rates of COVID-19 during our analysis period. Model findings were relatively unaffected by missing data assumptions in our sensitivity models (Table C, available as a supplement to the online version of this article at http://www.ajph.org).

TABLE 1— Characteristics of Wisconsin K–12 School Districts and Their Use of COVID-19 Prevention Strategies During the Fall 2021 Academic Semester: Wisconsin, September 2–November 24, 2021

	Distancing Policy (n = 278)			Masking Policy (n = 298)			Quarantine Policy (n = 169)		
	Robust, No. (%) or Mean (SD)	Absent, No. (%) or Mean (SD)	P	Robust, No. (%) or Mean (SD)	Absent, No. (%) or Mean (SD)	P	Robust, No. (%) or Mean (SD)	Absent, No. (%) or Mean (SD)	P
Districts[a]	188 (67.6)	90 (32.4)	. . .	73 (24.5)	202 (67.8)	. . .	87 (51.5)	79 (46.7)	. . .
% of district population fully vaccinated against COVID-19	52.8 ± 10.7	49.3 ± 9.2	.009	59.4 ± 10.9	48.7 ± 8.6	< .001	52.1 ± 10.5	49.8 ± 9.3	.49
Teacher age, y	44.6 ± 2.5	44.9 ± 2.1	.45	44.4 ± 2.4	44.6 ± 2.4	.3	45 ± 2.4	44.6 ± 2.0	.07
% of teachers vaccinated	77.6 ± 8.7	74.8 ± 10.7	.04	80.7 ± 8.2	75.6 ± 9.3	< .001	76.1 ± 10.8	75.7 ± 8.2	.6
Teachers per district	129.7 ± 190.8	123.2 ± 132	.4	215.2 ± 277.5	99.6 ± 106.4	.001	100.8 ± 95.8	122.9 ± 133.6	.51
Students per district	1827.7 ± 2760.7	1811.3 ± 2013.4	.29	3060.7 ± 3984.3	1428.7 ± 1636.9	.003	1427.7 ± 1442.7	1787.8 ± 2037.1	.48
Student:teacher ratio	13.2 ± 2.0	13.8 ± 2.1	.03	13.4 ± 2.1	13.4 ± 1.9	.98	13.2 ± 2.1	13.6 ± 1.9	.42
Use of policy for both students and staff[b]									
Districts with a robust distancing policy	188 (100)	0 (0)	. . .	63 (94)	103 (56.9)	< .001	58 (69)	38 (50)	.05
Districts with a robust masking policy	63 (34.8)	4 (4.5)	< .001	73 (100)	0 (0)	. . .	27 (31.8)	6 (7.8)	.005
Districts with a robust quarantine policy	58 (59.2)	26 (40)	.05	27 (79.4)	51 (43.6)	.005	87 (100)	0 (0)	. . .
Regular staff testing for COVID-19	20 (11.8)	11 (13.1)	.92	12 (18.8)	18 (9.9)	.17	10 (12.2)	4 (5.6)	.15

Note. The study involved n = 307 school districts.

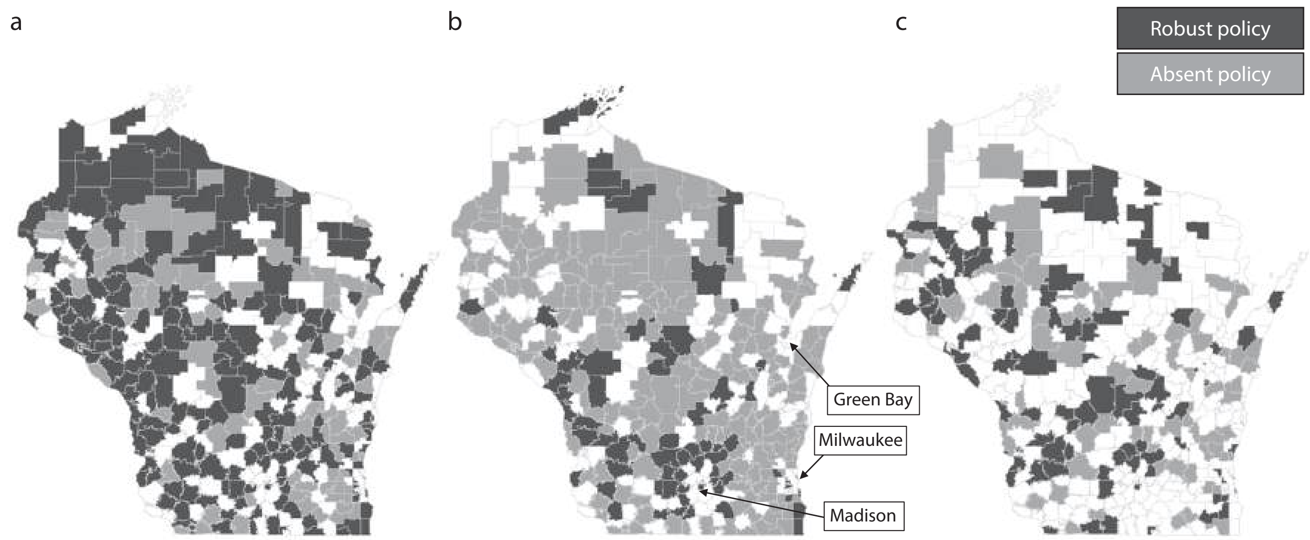
[a]Robust indicates policy in place for both students and staff. Absent indicates policy in place for neither students nor staff. The partial categories of any given policy (i.e., policy in place for either students or staff) are not presented because of small numbers. The total number of districts with a response for given policy is used as the denominator used for percentage calculations.
[b]Total districts with given policy implementation is used as the denominator used for percentage calculations.

When we stratified our imputed model by grade level, the hazards reduction associated with a robust masking policy remained consistent and statistically significant across elementary, middle, and high school locations (HR = 0.83 [CI = 0.77, 0.99]; HR = 0.74 [CI = 0.58, 0.95]; and HR = 0.77 [CI = 0.61, 0.98], respectively).

In assessing the potential for outcome ascertainment bias among school districts, we noted that the use of COVID-19 testing policies among educators was low but comparable between districts using different COVID-19 prevention policies (Table 1). In addition to unadjusted Kaplan–Meier curves (Figure B, available as a supplement to the online version of this article at http://www.ajph.org), we also reran our complete case model (Table 2; model 1), including a binary indicator variable for staff testing alongside the 3 other policy variables; it did not substantially alter the point estimates or CIs for our main policies of interest (not shown).

DISCUSSION

Our results provide further evidence of the benefits of student and staff masking in school settings during a period of high community transmission.[1–3,25,26] COVID-19 incidence rates in our assessed group of Wisconsin K–12 school district communities averaged 49.3 per 100 000 residents during the study period (range = 2.6–293.6 per 100 000 residents). During the first 3 months of the 2021–2022 academic year (September 2–November 24), and adjusted for pertinent person- and community-level factors, in-person educators working in school districts with both student and staff masking policies in place were 19% less likely to experience a COVID-19 illness than



(total responses = 278)

Robust:	188/278	(67.6%)
Partial:	0/278	(0%)
Absent:	90/278	(32.4%)
Missing:	36/307	(11.7%)

(total responses = 298)

Robust:	73/298	(24.5%)
Partial:	23/298	(7.7%)
Absent:	202/298	(67.8%)
Missing:	9/307	(2.9%)

(total responses = 169)

Robust:	87/169	(51.5%)
Partial:	3/169	(1.8%)
Absent:	79/169	(46.7%)
Missing:	138/307	(45.0%)

FIGURE 2— Wisconsin K-12 Public School Districts' Implementation of COVID-19 Prevention Policies of (a) Distancing, (b) Masking, and (c) Quarantine: Fall 2021 Academic Semester

Note. Robust policies indicate those applied to both students and educators. Absent policies were not required for either students or educators. Because of small numbers, districts with partial policies (or those differentially applied to students and staff) are not highlighted in a color, nor are districts with missing policy data.

AJPH December 2022, Vol 112, No. 12

were their counterparts working in districts without any masking policies. This is equivalent to a 23% higher HR among educators in districts without masking policies than among educators in districts with robust masking policies. Moreover, the protective effect associated with a robust masking policy was consistent across elementary, middle, and high school environments.

Our findings complement those of an ecologic study by Budzyn et al., who, using the same MCH survey policy data, determined that after the start of the 2021–2022 school year, US counties with school mask policies in place for students and staff experienced a significantly lower risk of pediatric COVID-19 cases than did counties without mask mandates (16.3 vs 34.9 cases per 100 000 children aged < 18 years).[27] Existing research also corroborates our study's lack of association between COVID-19 risk and either physical distancing or quarantine after exposure (imputed HR = 1.08; 95% CI = 0.98, 1.19 and HR = 0.98; CI = 0.89, 1.07, respectively). For example, 2 articles from the Duke University School of Medicine that suggest that—in the presence of masking policies—distancing or quarantine policies might have little effect on COVID-19 risk reduction.[25,26]

K–12 educators, despite a higher risk of workplace-associated COVID-19 incidence, do not appear to be at more risk for severe outcomes of COVID-19 than do those in other professional categories.[28,29] But, in our work, the 23% higher rate of COVID-19 illness among educators in districts without any masking policy is not without potential ramifications. In studies of school-based COVID-19 outbreaks, researchers identified that staff are often as involved in outbreaks as students.[4,6,12,13] These school-based outbreaks can subsequently spill over to the surrounding community members; for instance, preventing COVID-19 transmission in educational settings has a noted benefit to households associated with schoolchildren.[15]

We also note that the educators in our study were relatively young (average age = 44 years), almost entirely non-Hispanic White (97.1%), and highly vaccinated (77.9% having completed a full, primary vaccination series by the start of school). Therefore, our calculated HRs among Wisconsin educators might not be generalizable to all educators in the United States. Indeed the 23% higher HR of COVID-19 associated with a lack of a masking policy in Wisconsin school

TABLE 2— Effect of School District Policy (Physical Distancing, Masking, and Quarantine) on Hazard Rate of COVID-19 Among K–12 Educators, Stratified by Grade Level: Wisconsin, September 2–November 24, 2021

School Setting and Policy	Model Data Assumptions[a] HR (95% CI)[b]	
	Complete Cases	Imputed
Elementary school		
Distancing	1.14 (0.93, 1.40)	1.06 (0.92, 1.23)
Masking	0.78 (0.60, 1.03)	0.83 (0.70, 0.99)
Quarantine	1.01 (0.82, 1.23)	1.06 (0.92, 1.21)
Middle School		
Distancing	1.05 (0.76, 1.47)	0.97 (0.79, 1.19)
Masking	1.02 (0.68, 1.54)	0.74 (0.58, 0.95)
Quarantine	0.89 (0.66, 1.22)	0.99 (0.81, 1.20)
High School		
Distancing	1.23 (0.96, 1.60)	1.19 (0.99, 1.44)
Masking	0.76 (0.53, 1.07)	0.77 (0.61, 0.98)
Quarantine	0.96 (0.75, 1.23)	0.91 (0.76, 1.08)
Overall		
Distancing	1.11 (0.96, 1.28)	1.08 (0.98, 1.19)
Masking	0.81 (0.67, 0.98)	0.81 (0.72, 0.92)
Quarantine	1.00 (0.87, 1.15)	0.98 (0.89, 1.07)

Note. CI = confidence interval; HR = hazard ratio.

[a]Reflects 2 data sets treating missing policy data in distinct ways: (1) only school districts with complete data for all 3 policies or (2) imputed data for missing policy information using information from nonmissing district-level characteristics. Multivariate model adjusted for each of the 3 policies (masking, distancing, quarantine), teacher full vaccination status by start of school, National Center for Education Statistics school district locale, and spline terms for the following variables: teacher age (years), percentage of school district community fully vaccinated, weekly COVID-19 incidence rate in the school district community, and average student:teacher ratio in school district.
[b]HRs and CIs associated with policy implementation, robust vs absent. Robust indicates policy in place for both students and staff. Absent indicates policy in place for neither students nor staff. The partial categories of any given policy (i.e., policy in place for either students or staff) are not presented because of small numbers.

districts could be more pronounced in US school districts with an older or less vaccinated population of educators.

Limitations and Strengths

The findings of this study are subject to at least 3 principal limitations. First, policy variables were based on responses at the beginning of the semester. We were unable to account for potential changes to policy throughout the semester. However, we note that the trajectory of COVID-19 cases in Wisconsin was increasing from early July 2021 through mid-January 2022. For this reason, we do not expect that policies were suspended during our analysis period—if anything, it is more likely that some districts without policies in September implemented them during the analysis period. In this sense, our results might reflect conservative estimates. Similarly, although we were unable to account for measures of policy compliance, we do not anticipate that policy compliance dramatically waned during this period of increasing case rates—at least not because of a lack of pandemic awareness throughout the state.

Second, the MCH survey requested answers to broad questions (Table A, available as a supplement to the online version of this article at http://www.ajph.org). Because of this, the categorical exposure levels in our analysis might obscure nuances in the way distancing, masking, or quarantine policies were implemented in each district or among schools in the district. For example, there were no data available regarding the type of masks required in school districts with masking policies. Evidence shows that different types of masks are associated with different levels of fit, quality, and effectiveness,[30–32] and so our overall risk reduction associated with masking may gloss over more nuanced levels of protection associated with various masks.

Similarly, because of small numbers, we were unable to assess risks of COVID-19 associated with a heterogeneous application of policies, such as the effects of staff masking or student masking alone. We cannot conclude, therefore, whether mask wearing by in-person educators or by students specifically contributed more to the reduction in educator risk. Future work could consider the risk reduction in schools with a mask policy applied only to in-person educators.

Third, there was potential for selection bias in our analysis, although we took care to minimize any potential consequences of this. It is true that, statewide, 11% of all regular K–12 school districts did not report any policy data, and we excluded these from analysis. However, these districts were distributed throughout the state in urban and rural areas, which minimized the concern of unrepresentative data (Figure 2). Similarly, it is possible that educators in different districts were more or less likely to report COVID-19 cases to local health departments, perhaps because of prevailing social willingness to be tested for COVID-19 or the use of self-tests at home. To lessen the impact of this bias, we included a random effect term for school district in our model.

This study builds on the existing literature in 2 notable ways. For one, previous studies investigating COVID-19 prevention policies in schools often lacked comparison groups because of their analysis time frame, which occurred when the vast majority of school districts had implemented similar masking and other prevention policies; these previous studies were limited in ability to contrast policies. Previous studies often considered only schools in which the policy was applied, and thus researchers were unable to determine whether the observed low COVID-19 risk was associated with the presence of the prevention policy itself. In our analysis of heterogenous policy use, we found that the presence of student and staff masking policies in Wisconsin school districts, compared with the absence of such policies, was associated with a significantly reduced rate of COVID-19 among in-person educators.

A second strength of our analysis was our ability to control for a wide range of pertinent person- and community-level confounders. We were able to use data from a variety of state and national data sources to control for educator vaccination status, educator age, community vaccination status, weekly incidence of COVID-19 in the community, urbanicity of the school district, and student to teacher ratio. Additionally, we implemented a random-effects model in an attempt to control for unobserved confounders at the school district level.

Public Health Implications

Our work shows that an in-person educator's risk of infection can be reduced with group mask use—a simple, nonpharmaceutical intervention. Beginning in February 2022, the Omicron variant wave of the COVID-19 pandemic tapered off, prompting the United States and other countries to lift many or all of their societal COVID-19 prevention policies. Fortunately, surveillance data continue to indicate that the risk of severe COVID-19 outcomes in younger children remains rare. But in considering the beneficiaries of masking policies in US K–12 schools, it is important to bear in mind the health of the nation's 5.5 million K–12 educators and the 3 million additional in-school staff.[14]

We want to be clear that our findings do not suggest that a robust mask policy in K–12 schools be applied in perpetuity without consideration of external factors. Instead, our work adds further evidence to underscore the role of mask policies in school environments. Student and staff mask wearing during periods of high community transmission prevented illness in schools among a highly vaccinated population of in-person educators and may be a worthwhile consideration during future periods of high COVID-19 transmission in the community. AJPH

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CONTRIBUTORS

P. M. DeJonge designed the study, led statistical analyses, and drafted the article. P. M. DeJonge, C. Tomasallo, and J. Meiman codeveloped the analysis plan. I. W. Pray and K. McCoy provided epidemiologic and subject matter expertise. R. Gangnon provided statistical expertise and reviewed methods for validity throughout the analysis. C. Tomasallo and J. Meiman provided major contributions in database access and management at the state level. All authors reviewed and revised the final version of the article.

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Note. The findings and conclusions in this report are those of the authors and do not necessarily represent the official position of the Centers for Disease Control and Prevention (CDC).

CONFLICTS OF INTEREST

The authors have no potential or actual conflicts of interest from funding or affiliation-related activities to disclose.

HUMAN PARTICIPANT PROTECTION

The CDC reviewed this activity and determined that it met the requirements of public health surveillance as defined in 45 CFR 46.102(l)(2) and was conducted in a manner consistent with applicable federal law and CDC policy. We stored all identifiable information on protected Wisconsin Department of Health Services servers, and analysts used only de-identified data.

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Excess Mortality From Non–COVID-19 Causes During the COVID-19 Pandemic in Philadelphia, Pennsylvania, 2020–2021

Megan Todd, PhD, and Annaka Scheeres, MPH

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Objectives. To estimate excess mortality from non–COVID-19 causes during the COVID-19 pandemic in Philadelphia, Pennsylvania, and understand disparities by race/ethnicity, age, and sex.

Methods. We used Poisson regression models of weekly deaths using data from Pennsylvania's vital registration system (2018–2021).

Results. There was significant excess mortality as a result of heart disease, homicide, diabetes, drug overdoses, traffic crashes, and falls in 2020–2021; the burden of this excess non–COVID-19 mortality fell on non-Hispanic Black Philadelphians. Among younger non-Hispanic Black men, homicide and drug overdoses were responsible for 54% and 18% of excess deaths—more than COVID-19 (17%). For younger non-Hispanic Black women, drug overdoses accounted for 51% of excess deaths, whereas COVID-19 accounted for 40%.

Conclusions. Excess mortality was not solely caused by severe acute respiratory syndrome coronavirus 2 (SARS-CoV-2; the causative agent of COVID-19), particularly at younger ages. Indirect pandemic mortality exacerbated prepandemic disparities by race/ethnicity.

Public Health Implications. Excess mortality as a result of non–COVID-19 causes may reflect indirect pandemic mortality. National cause-of-death data lag behind local cause-of-death data; local data should be examined as an early indication of trends and disparities. Public health practitioners must center health equity in pandemic response and planning. (*Am J Public Health*. 2022;112(12):1800–1803. https://doi.org/10.2105/AJPH.2022.307096)

The COVID-19 pandemic caused a dramatic increase in mortality, but not all of this excess mortality is directly attributable to infection with severe acute respiratory syndrome coronavirus 2 (SARS-CoV-2; the causative agent of COVID-19).[1] The pandemic caused profound disruptions in society, which may have led to excess mortality indirectly related to the virus. Researchers have speculated about these indirect pathways—such as interruptions in health care[2,3] and worsening mental health[4]—but so far, little work has studied excess mortality as a result of non–COVID-19 causes.

In this study, we estimated excess mortality as a result of non–COVID-19 causes of death in Philadelphia, Pennsylvania. Past studies have documented differences in COVID-19 mortality by sex,[5] age,[6] and race/ethnicity[7,8]; we therefore compared mortality by these demographic characteristics to see if this was also the case for non–COVID-19 mortality. National cause-of-death data lag behind local cause-of-death data; these data from Philadelphia—the sixth largest US city—provide a timely estimate of trends and disparities in mortality for 2020–2021.

METHODS

Data are from Pennsylvania's vital registration system. We used final 2018–2019 death files, combined with preliminary 2020–2021 files (updated June 30, 2022), to examine mortality in Philadelphia from January 1, 2018, to January 1, 2022. Deaths are reported with a delay; for more details, see section 1 of Appendix (available as a supplement to the online version of this article at http://www.ajph.org). We calculated excess mortality rates with denominators from the US Census Bureau's 2021 Annual County Resident Population Estimates.

Following Todd et al.,[9] we trained Poisson models of weekly mortality on 2018–2019 data, stratified by age, sex, and race/ethnicity and allowing for seasonal trends. Our past work examined all-cause mortality through 2020[9]; here, we added cause-specific mortality from the most common pre-COVID-19 causes of death (heart disease, cancer, injury [disaggregated into homicide, drug overdoses, traffic crashes, and falls], cerebrovascular disease, diabetes, septicemia, influenza and pneumonia, chronic respiratory diseases, and chronic kidney diseases) and data through 2021. (See section 2 of online Appendix for model details.) We then used these models to estimate expected cause-specific mortality from March 15, 2020, to January 1, 2022 by sex (male, female), age group (< 50 years old, ≥ 50 years), and race/ethnicity (non-Hispanic Black, non-Hispanic White; other categories omitted because of small counts). We compared expectations with observed deaths to obtain estimates of cause-specific excess mortality. All deaths as a result of COVID-19 were considered excess deaths. We conducted the analysis using R 4.1.1 (R Foundation for Statistical Computing, Vienna, Austria).

RESULTS

There were 5963 excess deaths from all causes between March 15, 2020, and January 1, 2022, representing 23% more deaths than predicted (Table A, available as a supplement to the online version of this article at http://www.ajph.org); 4469 (75%) of these excess deaths were directly attributable to COVID-19. Among non–COVID-19 causes of death, the greatest proportional increases above expectations occurred for deaths caused by traffic crashes (53% more deaths than expected), homicide (51%), and diabetes (41%). Significant increases above expectations were also observed for deaths caused by falls (22%), drug overdoses (16%), and heart disease (6%). As the most common cause of death, heart disease was responsible for the largest number of excess deaths (n = 375) of any non–COVID-19 cause, despite only a modest percentage increase. Homicide was responsible for the second largest number of excess deaths (n = 327), followed by drug overdose (n = 272) and diabetes (n = 244). As less common causes of death, traffic crashes and falls accounted for 95 and 53 excess deaths, respectively. There was a decrease below expectations for deaths from chronic respiratory diseases (9%, or 89 fewer than expected). Observed deaths were not significantly different from expectations for cancer, kidney disease, pneumonia and influenza, stroke, or septicemia.

Excess mortality was not distributed equally; the burden fell more heavily on non-Hispanic Black Philadelphians than non-Hispanic White Philadelphians. Figure A (available as a supplement to the online version of this article at http://www.ajph.org) shows excess deaths per 100 000 from all causes by sex, age group, and race/ethnicity. There was significant excess mortality among adults aged 50 years and older in all sex–race groups. However, at younger ages, only non-Hispanic Black men and women experienced excess mortality (241 and 70 excess deaths per 100 000, respectively), whereas non-Hispanic White women and men did not experience significant excess mortality.

Figure 1 disaggregates excess mortality for age–sex–race/ethnicity groups by cause. Only causes of death for which the number of deaths was significantly different from expectations are labeled; see Table B (available as a supplement to the online version of this article at http://www.ajph.org) for complete counts. For those aged 50 years and older, COVID-19 was overwhelmingly responsible for excess mortality: there were 1142 COVID-19 deaths per 100 000 for older non-Hispanic Black men (representing 66% of excess deaths for this group), 859 per 100 000 for older non-Hispanic Black women (79% of excess deaths), 855 per 100 000 for older non-Hispanic White men (97%), and 661 per 100 000 for older non-Hispanic White women (over 100%, a figure that might be attributable to declines from other causes). For older non-Hispanic Black men, there was significant excess mortality from heart disease, drug overdoses, diabetes, and traffic crashes. Among older non-Hispanic Black women, mortality from heart disease and drug overdoses significantly exceeded expectations. For older non-Hispanic White men, diabetes was the only significant non–COVID-19 contribution to excess mortality. Older non-Hispanic White women also experienced significant excess

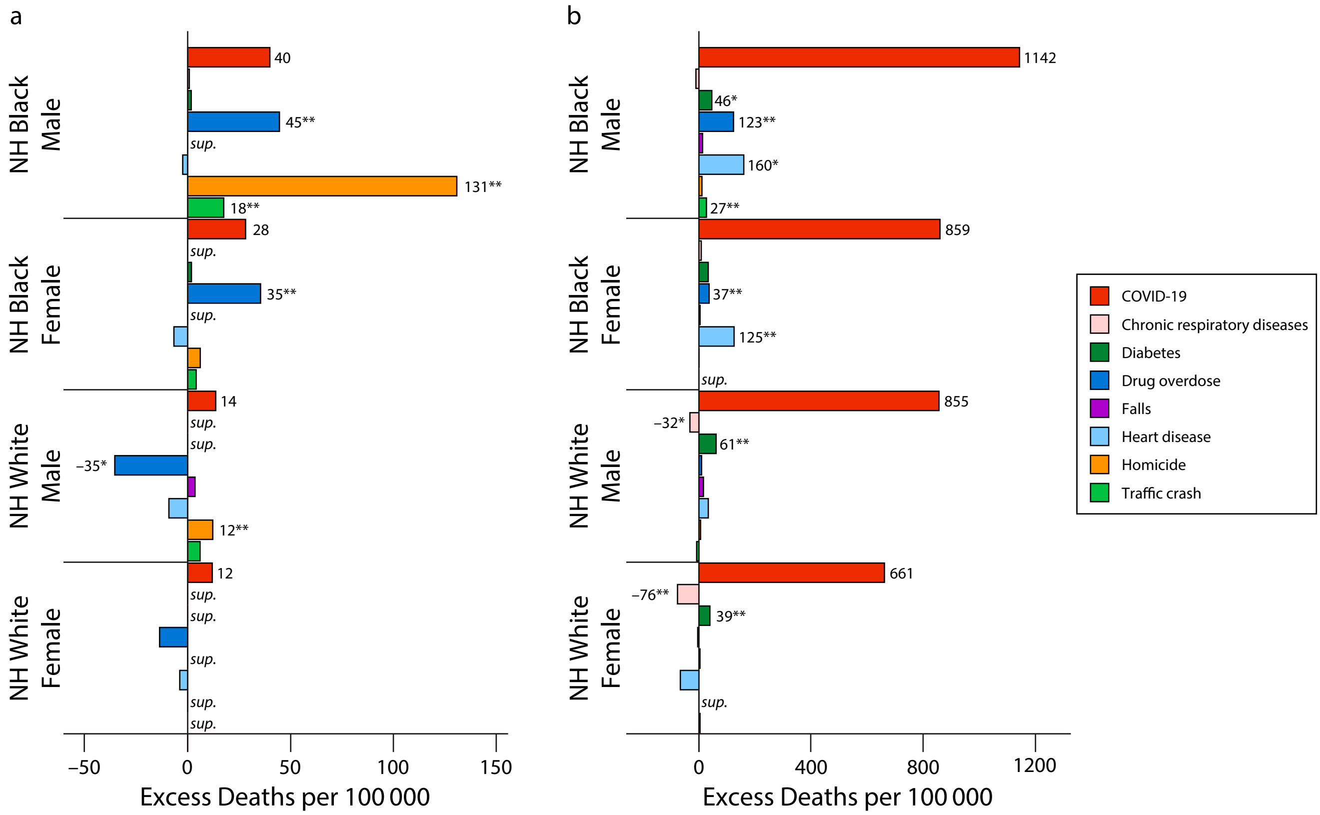

FIGURE 1— Excess Mortality per 100 000 Population by Cause of Death, Race/Ethnicity, and Sex for Those Aged (a) Younger than 50 Years and (b) 50 Years or Older: Philadelphia, PA, March 15, 2020–January 1, 2022

Note. NH = non-Hispanic; "sup." indicates suppressed value because observed count was less than 10. We calculated COVID-19 mortality rates assuming that (expected deaths) = 0.
**P* < .05, ***P* < .01 for test of the null hypothesis that (observed deaths) = (expected deaths).

mortality from diabetes, but this was more than offset by significant reductions in mortality from chronic respiratory diseases, possibly because of COVID-19 mitigation strategies like social distancing and masking.

For Philadelphians aged younger than 50 years, the contribution of COVID-19 to excess mortality was far more modest. Among young non-Hispanic Black men, COVID-19 was only the third leading cause of excess mortality (40 excess deaths per 100 000; 17% of excess deaths), trailing behind homicide (131 per 100 000; 54%) and drug overdoses (45 per 100 000; 18%). Traffic crashes also significantly contributed to excess mortality for young non-Hispanic Black men. For young non-Hispanic Black women, drug overdoses contributed more to excess mortality (35 per 100 000; 51% of excess deaths) than COVID-19 (28 per 100 000; 40%). There was no significant all-cause excess mortality among young non-Hispanic White men and women; excess COVID-19 deaths were offset by lower-than-expected mortality from drug overdoses and heart disease.

DISCUSSION

This study estimated cause-specific excess mortality during the COVID-19 pandemic in Philadelphia. In addition to deaths from COVID-19, there was significant excess mortality from heart disease, homicide, diabetes, drug overdoses, traffic crashes, and falls. The burden of non–COVID-19 mortality disproportionately affected older non-Hispanic Black Philadelphians compared with older non-Hispanic White Philadelphians. Among younger non-Hispanic Black Philadelphians, COVID-19 mortality was dwarfed by excess mortality from homicide and drug overdoses. Excess non–COVID-19 mortality may have resulted from interruptions in health care (for heart disease, diabetes, and drug overdoses), or from stress, anxiety, and mental strain (all causes). Although the number of traffic crashes in Philadelphia decreased in the first year of the pandemic, the number of fatalities increased, possibly because of excess speed amid reduced traffic volume.[10] Although more research is needed to understand why non–COVID-19 causes of death contributed to excess mortality during the pandemic, our work shows that this excess mortality was substantial, and contributed to mortality disparities by race/ethnicity.

PUBLIC HEALTH IMPLICATIONS

Preexisting racial mortality disparities were exacerbated by COVID-19.[7] This study is preliminary evidence that non–COVID-19 mortality during the pandemic further contributed to disparities, notably at younger ages, where the mortality risk from COVID-19 was small. This is an urgent call to think broadly about the impacts of COVID-19 on health and mortality and to center equity in pandemic response and preparedness planning. AJPH

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CONTRIBUTORS

M. Todd originated and led the study. A. Scheeres conducted analyses. Both authors contributed to writing.

CONFLICTS OF INTEREST

The authors have no conflicts of interest to disclose.

HUMAN PARTICIPANT PROTECTION

This study was determined to be exempt by the Philadelphia Department of Public Health institutional review board.

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Considering Potential Risks Associated with Coopetition in Social Determinants of Health

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Butler and Nichols[1] highlight the possibilities of coopetition (i.e., cooperative competition at interorganizational and intraorganizational levels) to fund the infrastructure of social determinants of health (SDoH). Although early examples showed success, coopetition poses possible risks to community-based organizations (CBOs) that offer SDoH services.

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RISK 1: COST

A formal coopetition mechanism can be expensive and may impose above average functioning costs on participating organizations.[2] Increased functioning costs may reduce the overall philanthropic efforts in CBOs outside the coopetition model. Thus, it is important to evaluate the total cost and total income at the CBO level before and after coopetition.

RISK 2: EQUITABLE FUNDING AND PURCHASING

Health plans possess tremendous financial power and can influence the purchasing decisions of other funders in coopetition. For instance, health plans are innately more interested in SDoH solutions that generate short-term, clear returns on investment for them (e.g., food) than in other solutions (e.g., home modifications, family caregiver support).[3] It is possible that solutions with less return on investment evidence experience reduced funding. Tracking the funding status of services that are outside the coopetition model at a community level will offer a more comprehensive picture of coopetition's impact.

RISK 3: AUTONOMY

Coopetition often requires a CBO network lead to negotiate on behalf of a group of CBOs. Although the CBO network lead plays an important role, some emerging evidence in coopetition shows that formal hierarchical structure has a negative effect on knowledge sharing, whereas informal lateral relations (e.g., social interactions) have a positive effect.[4] In particular, a hierarchical model may unintentionally harm knowledge sharing and capacity building for small, minority-led organizations. Coopetition models should maximize autonomy and lateral interactions.

RISK 4: RESEARCH AND DEVELOPMENT

Radical innovation in SDoH is needed. One known advantage of coopetition is accelerating research and development. Yet coopetition in SDoH today still focuses too much on providing SDoH services and information exchanges, which may limit flexibility in research and development.[5] Two practices may catalyze research and development. First, long-term SDoH coopetition is encouraged because coopetition that spans five to seven years is more likely to generate benefits related to increased innovation.[2] Second, coopetition should treat CBO-led research and development as part of the infrastructure and allow flexible funding for such activities.

The US social care system is at a tipping point. Thoughtful coopetition that prioritizes structural, long-term benefits for CBOs warrants further research. AJPH

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CONFLICTS OF INTEREST

V. Wong is one of the founders of EVISET, a new venture aiming to improve structural equity in the field of social care.

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Butler and Nichols Respond

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We are grateful for Wong's acknowledgment of the benefits of health plan co-opetition in building and operating the infrastructure for emerging health and social service ecosystems. She is correct that there are risks. But the gap between investment in social determinants of health (SDOH) and unmet human need is so great that the benefits of encouraging collaborative investment in SDOH infrastructure and services outweigh those risks.

Wong's main point is that the interests of community-based organizations (CBOs) and their clients must be prioritized. We not only agree; we argue that health plan co-opetition is more likely to produce this result than other arrangements.

We emphasized that a well-functioning ecosystem requires trust, technology, data management, contracting expertise, and sustainable financing. We agree that trust among CBOs and clients who need services is key. But the other infrastructure elements are also essential. Our point is that health plan collaboration is more efficient than hyper-competition in which plans contract for everything on their own, raising costs for plans and CBOs alike through redundant technologies and contracting and data reporting requirements. Plans now engaged in collaboration (and even some that are not) recognize the importance of community trust in their own success; this will fuel alignment over time.

Wong views network lead entities, which contract and manage data on behalf of CBOs, as a risk to knowledge sharing and capacity building within CBOs themselves. We agree that network lead entities cannot function well without the trust of CBOs and their clients; therefore, they must emerge organically and locally and work to expand CBO capacities. Still, many CBOs today lack the contracting and data management capacity essential for value-based contracting, so the functions of network lead entities are catalytic for the spread of cost-effective SDOH investments.

Some health plan–led SDOH investment may indeed focus mainly on reducing short-run costs, but our CommonSpirit and CAPGI (Collaborative Approach to Public Goods Investments) examples and plan investments in housing suggest that innovation and longer-run, more social return on investment are the foci of engaged plans. Moreover, the broader social benefits of this form of co-opetition are attractive to government policymakers. Health care is learning that many clients have more than one unmet need and that leaving social needs unmet reduces the effectiveness of each intervention. Although health organizations cannot pay for utopia alone, health plan co-opetition is a promising tool, perhaps in partnership with government investment, for expanding commitments to social needs and the infrastructure on which effective delivery depends. AJPH

AJPH Peer Reviewers: Thank You for Your Services

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