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Short Communication

COVID-19 preventive behaviours among people with anxiety and depressive symptoms: findings from Japan

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ABSTRACT

Objectives: The aim of the study was to examine COVID-19 preventive behaviours among individuals with mental health problems.
Study design: This is a pooled cross-sectional study.
Methods: Online survey data were analysed from 2000 Japanese adults collected in April and May 2020. Information was obtained on 13 COVID-19 preventive behaviours and anxiety and depressive symptoms using the Generalized Anxiety Disorder 7-item scale and Patient Health Questionnaire-9, respectively. Linear regression analysis was used to examine the associations.
Results: In models adjusted for demographic and socio-economic factors, anxiety (coefficient: −0.77, 95% confidence interval [CI]: −1.30, -0.24) and depressive symptoms (coefficient: −0.82, 95% CI: −1.34, -0.30) were both associated with significantly lower engagement in COVID-19 preventive behaviours.
Conclusion: Our results highlight the importance of facilitating the performance of preventive behaviours in individuals with mental health problems to prevent the spread of COVID-19 in this population.

As yet, there has been little research on the association between common mental disorders (CMDs) and infectious disease, and it is thus uncertain how the presence of pre-existing CMDs such as anxiety and depression might be linked to health behaviours during the ongoing COVID-19 pandemic. A recent editorial has suggested, for example, that differing levels of health anxiety might result in either a maladaptive engagement in (e.g., excessive handwashing [high anxiety]) or a disregard for (e.g., no handwashing [low anxiety]) preventive behaviours.[1] In addition, other authors have previously hypothesised that characteristics associated with CMDs such as depression, including lower levels of energy, a decreased focus, and greater hopelessness, might also be important for (non-)engagement in necessary health behaviours.[2]

The few studies that have examined how CMDs are linked to infectious disease preventive behaviours have produced mixed results. An earlier study from Hong Kong found that compared with those with low anxiety, individuals with high and, especially, moderate levels of anxiety had significantly higher odds for adopting ≥5 precautionary measures against severe acute respiratory syndrome.[3] Support for the notion that CMDs might be associated with increased engagement in preventive behaviours also comes from another study from Hong Kong, which recently found that people with symptoms of anxiety were more likely to adopt social distancing measures in response to the threat of COVID-19.[4] In contrast, other research from China found that anxiety was not related to any differences in the adoption of preventive measures, while people with depression took fewer preventive measures in response to the COVID-19 pandemic.[5]

The present study will examine the effects of anxiety and depressive symptoms on COVID-19 preventive behaviours in a sample of the Japanese general population. A focus on Japan may be particularly instructive. Although the effects of COVID-19 have not been as severe in Japan as in many other countries—at least in terms of the number of deaths—coronavirus cases began to increase quickly from early July to mid-July after the ending of a nationwide state of emergency in late May. This increase may be linked to several factors including the use/non-use of preventive measures. Specifically, a recent study has reported that although the vast majority of

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Japanese adults have adopted preventive measures, around 20% of the working-age population (age: 20–64 years) is reluctant to do so.[6] However, that study did not specifically focus on the possible effects of CMDs on the non-use of preventive behaviours.

Two rounds of an online survey of the Japanese population were administered between April 16 and April 18, 2020 (1st round), and between May 15 and May 17 (2nd round). A commercial survey company, the Survey Research Center, was tasked to send out a set of screening questions to approximately 10,000 respondents from its commercial web panel and then to construct a sample of 1000 respondents based on their demographic characteristics in each round. A new set of respondents was drawn in the second round. The final sample comprised respondents who were representative of the Japanese general population in terms of the area of their residency, sex, and age distribution. The respondents in the final sample answered online questions about their mental health, personal economic situation, and preventive behaviour with regard to COVID-19, among others. The final sample size was 2000. The self-report Patient Health Questionnaire-9 was used to assess depressive symptoms in the past two weeks.[7] A score of 10 or higher (out of 27) was regarded as a case of at least moderate depressive symptomatology. The Cronbach's alpha value for the scale was 0.90. The self-administered Generalized Anxiety Disorder 7-item scale was used to measure anxiety symptoms in the past two weeks.[8] A score of 10 or higher (out of 21) was regarded as a case of at least moderate anxiety. The scale had a good degree of internal reliability (Cronbach's alpha = 0.92). Information was collected on 13 COVID-19 preventive behaviours (no/yes). The specific behaviours and their frequency are detailed in Appendix A in the supplementary data. Information was also obtained on a number of covariates including age, sex, education, income, household financial situation (versus previous year), employment status, and data survey round.

Linear and logistic regression analyses were performed to examine the association between anxiety and depressive symptoms and preventive behaviours. In the first analysis, a combined preventive behaviour score variable was created by summing the responses for each preventive behaviour, and linear regression analysis was used to examine the associations. In the second analysis, the association between anxiety and depressive symptoms and each of the individual preventive behaviours was examined using binomial logistic regression. All analyses were adjusted for the previously listed covariates. The standard errors were heteroskedasticity robust and clustered by prefecture. The analysis was conducted using STATA/MP (version 16, Stata Corporation, College Station, TX). The results are presented as coefficients (Coef.) and odds ratios (ORs) with 95% confidence intervals (CIs). The level of statistical significance was set at $P < 0.05$ (two-tailed).

The frequency of anxiety and depressive symptoms was 10.9% and 17.3%, respectively. Both anxiety (Coef: −0.77, 95% CI: −1.30, -0.24) and depression (Coef: −0.82, 95% CI: −1.34, -0.30) were associated with significantly reduced engagement in all of the preventive behaviours combined (Table 1). For anxiety, in the logistic regression analyses, ORs were negative for 11 of the 13 preventive behaviours (Appendix A). Individuals with symptoms of anxiety were significantly less likely to engage in six of the preventive behaviours. Specifically, they had a 40–49% reduction in the odds for washing hands, wearing a mask, and avoiding crowds and a 26–38% reduction in the odds for using a tissue/sleeve when coughing/sneezing, avoiding touching the face, and cancelling going out. Depressive symptoms were also associated with significantly reduced odds for the same six preventive behaviours. In addition, they were also associated with a 27% reduction in the odds for avoiding engaging in gatherings (OR: 0.73, 95% CI: 0.58–0.90).

Although a study from China reported that neither state nor trait anxiety was associated with COVID-19 preventive behaviours,[5] other recent studies have all linked anxiety with an increased likelihood of engaging in preventive behaviours.[4,9,10] This conflicts with our finding that anxiety symptoms were associated with reduced preventive behaviour. It is uncertain what underlies this difference, but underlines the need for future studies to collect information on the specific causes of anxiety, especially as it has been suggested that high levels of 'health' anxiety might be linked to engaging in excessive preventive behaviour.[1] Regarding depression, our findings accord with those from the above-mentioned Chinese study, which showed that depressive symptoms may inhibit preventive behaviours in response to the COVID-19 pandemic.[5] It is possible that various mechanisms might underlie the association between CMDs and reduced preventive behaviour in Japanese adults. For example, it can be speculated that symptoms that are characteristic of these disorders such as fatigue and reduced concentration might be important in this regard.

Table 1
Association between anxiety and depressive symptoms and all COVID-19 preventive behaviours combined among Japanese adults.[a]

	Anxiety	Depression
	Coef. (95% CI)	Coef. (95% CI)
Preventive behaviours	−0.77 (−1.30, −0.24)[b]	−0.82 (−1.34, -0.30)[b]

Coef: coefficient; CI: confidence interval.
Both analyses were adjusted for age (reference: young), sex (reference: female), education (reference: less than college), income (reference: high income), household financial situation (reference: better/same as in the previous year), employment (reference: unemployed not in the labour force), and data survey round (reference: round 1).
[a] Anxiety and depressive symptoms were the exposures; COVID-19 preventive behaviours were the outcomes.
[b] $p < .01$.

This study has some limitations. The use of cross-sectional data meant that we were not able to establish causality or the direction of the observed associations. In addition, we also lacked information on prior psychiatric diagnoses of the respondents. It is possible, therefore, that poorer mental health might have been a psychological response to the threat of COVID-19 or the rigours of quarantine. Keeping this in mind, the results of this study indicate that people with mental health problems may be at increased risk of COVID-19 infection, given their lower engagement in a number of preventive behaviours. This highlights the importance of educating individuals with poorer mental health about the dangers of COVID-19 and how to protect themselves against the virus. In addition, our findings also suggest that further research on the effects of COVID-19 among individuals with mental health problems is now urgently warranted.

Author statements

Ethical approval

This study was approved by the Ethics Committee of Waseda University (approval case number: 2020-050) and Osaka School of International Public Policy, Osaka University. The survey participants were informed of the purpose of the study before their participation and had the option to quit the survey at any time. The respondents provided explicit consent that the information they provided could be used for the purpose of this study. The data are completely anonymous.

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Competing interests

The authors declare that they have no conflicting interests.

Author contributions

M.U. and T.M. were responsible for data acquisition. A.S. wrote the main text. M.U. analyzed the data and contributed to the writing of the text. T.M. and H.S. reviewed and revised the manuscript.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2020.09.017.

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Original Research

Hospital admission rates, length of stay, and in-hospital mortality for common acute care conditions in COVID-19 vs. pre-COVID-19 era

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ABSTRACT

Objectives: The impact of COVID-19 upon acute care admission rates and patterns are unknown. We sought to determine the change in rates and types of admissions to tertiary and specialty care hospitals in the COVID-19 era compared with pre-COVID-19 era.

Methods: Acute care admissions to the largest tertiary care referral hospital, designated national referral centers for cardiac, cancer and maternity hospital in the State of Qatar during March 2020 (COVID-19 era) and January 2020 and March 2019 (pre-COVID-19 era) were compared. We calculated total admissions, admissions for eight specific acute care conditions, in-hospital mortality rate, and length of stay at each hospital.

Results: A total of 18,889 hospital admissions were recorded. A sharp decline ranging from 9% to 75% was observed in overall admissions. A decline in both elective and non-elective surgeries was observed. A decline of 9%–58% was observed in admissions for acute appendicitis, acute coronary syndrome, stroke, bone fractures, cancer, and live births, whereas an increase in admissions due to respiratory tract infections was observed. Overall length of stay was shorter in the COVID-19 period possibly suggesting lesser overall disease severity, with no significant change in in-hospital mortality. Unadjusted mortality rate for Qatar showed marginal increase in the COVID-19 period.

Conclusions: We observed a sharp decline in acute care hospital admissions, with a significant decline in admissions due to seven out of eight acute care conditions. This decline was associated with a shorter length of stay but not associated with a change in in-hospital mortality rate.

Introduction

The ongoing COVID-19 pandemic has strained global healthcare capacity to breaking point in many countries.[1–4] Various approaches have been employed by different countries in response to the surge of cases, including physical distancing measures, targeted lockdowns, wider movement restrictions, travel bans, and closing of places of mass gatherings. Despite these measures, the number of persons with COVID-19 infection requiring hospitalization has exceeded acute and critical care bed capacity. To address acute and critical care bed capacity issues, most countries canceled or postponed elective surgeries and other admissions deemed non-vital to short-term patient outcomes. Early anecdotal reports in mainstream media noted a decrease in hospital admissions owing to 'heart attacks, strokes, and even appendicitis'.[5] This was confirmed by more recent reports noting a decline in admissions due to acute coronary syndrome (ACS), a decline in ST-segment elevation cardiac catheterization laboratory activations, a decrease in stroke imaging procedure performed, and an increase in out-of-hospital cardiac arrest.[6–9] An increase in emergency medical services arrival time and a decrease in by-stander initiated cardiopulmonary resuscitation were also noted.[7] The effect of COVID-19 pandemic

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upon admissions for general medical and surgical care, oncology services and obstetric services, and the effect of any change on in-hospital and overall community mortality rates is not known. Our aim was to determine the change in rates and reasons for hospitalization at four major referral centers in the State of Qatar. We chose the country's largest and main tertiary care hospital and dedicated hospitals and national referral centers for comprehensive cardiovascular care, cancer care, and obstetric and gynecologic services.

Methods

The study was conducted at Hamad Medical Corporation (HMC), the State owned and operated health-care system, which provides approximately 85% of acute care inpatient bed capacity in Qatar. The flagship and the largest hospital is Hamad General Hospital (HGH), which is the tertiary care referral center located in the capital city, Doha. Comprehensive cardiac care is provided at the Heart Hospital, which is the main designated facility in the country for primary percutaneous interventions for ACS. Cancer care is provided at the National Center for Cancer Care and Research (NCCCR), which is the only facility in the country providing comprehensive cancer care services, including chemotherapy, radiation therapy, and cyberknife services. Women's Wellness and Research Center (WWRC) is the largest provider of obstetric and gynecologic care in the country. These facilities were chosen because they represent the spectrum of acute tertiary and specialty care in the country. All hospitals use the same electronic health records system (Cerner®, Kansas City, MO, USA) which is interconnected, with patients retaining the same unique health identification number across all hospitals. All hospitals are accredited by The Joint Commission International and the clinical laboratories are accredited by the College of American Pathologists. All hospital discharges are reviewed by expert coders to assign the primary and secondary reasons for hospital admission. Up to ten reasons are recorded at the time of discharge using International Classification of Diseases, 10th edition Australian Modification.

The first case of COVID-19 in Qatar was diagnosed on February 28, 2020. As a matter of national policy, these four hospitals were designated to be COVID-19-free facilities. Any patient with known COVID-19 infection was admitted to a separate COVID-19 facility. Patients diagnosed incidentally or upon routine testing for COVID-19 after admission to any of these hospitals were evaluated and transferred to a designated COVID-19 facility, unless they were medically unstable for transfer. We determined the number of hospital discharges at each of the four hospitals for the month of March 2020 and compared it with January 2020 and March 2019. We omitted February 2020 because changes in hospital flow were being implemented in anticipation of the first wave of COVID-19 patients. All inpatient admissions at the four hospitals with a length of stay of more than one day were included. Data retrieved included patient demographics (age, sex, and nationality), admission and discharge dates and times, principal and up to nine additional discharge diagnoses, mode of arrival at the hospital, disposition, and surgical procedures performed (elective vs. non-elective).

We tabulated the overall number of admissions for each hospital across the three study periods, as well as the number of elective and non-elective surgical procedures performed, average daily admissions, average length of stay, and in-hospital mortality. Hospital use metrics, including mode of presentation to the hospital for that particular episode, and disposition were tabulated for each hospital, by study period.

Through consensus among coauthors, eight common diagnoses/conditions were chosen to represent the spectrum of conditions that generally require acute hospital care. These included acute appendicitis, ACS, other cardiovascular disease diagnoses (including cardiac arrhythmias, congestive heart failure, and angina pectoris without acute myocardial infarction), stroke, acute bone fractures, cancers, live births, and respiratory tract infections (excluding tuberculosis). Two study team members independently reviewed all diagnoses and identified and assigned them to one of the categories listed previously. These diagnostic groups were tabulated by study period, and the percentage change in each category was calculated.

To explore any correlation between acute care admission rates, in-hospital mortality and overall mortality rate in the State of Qatar, we retrieved publicly available data published by the Planning and Statistics Authority, State of Qatar, which publishes monthly reports on the total population and other vital statistics for persons residing in Qatar for that month. Number of deaths and the total population of Qatar for the months included in the study were retrieved to calculate the unadjusted death rate for each month.

Ethical approval

The study was reviewed and approved by the Institutional Review Board at Hamad Medical Corporation (MRC-05-034).

Results

A total of 18,889 hospital admissions were recorded for the three study periods (Table 1). There were 7545 admissions at HGH, 794 at HH, 396 at NCCCR, and 10,154 at the WWRC. Compared with January 2020, there was a 28.0% reduction in admissions at HGH, 59.6% reduction at HH, 45.7% reduction at NCCCR, and a 14.7% reduction at the WWRC. Compared with the same month in 2019, the reduction was 8.6% at HGH, 51.8% at HH, 75.5% at NCCCR, and a 21.5% at the WWRC. Number of elective surgeries dropped by 6.8% at the WWRC to 68.6% at HGH from January 2020 to March 2020, whereas non-elective surgeries showed a smaller decline. Average daily admissions and average length of stay also declined at all facilities over both time period comparisons (Table 1). In-hospital mortality rate was numerically lower at all hospitals for all time period comparisons except at HGH between March 2010 and January 2020. However, these comparisons were not statistically significant. There were no significant changes in terms of mode of presentation to the hospital or disposition at any of the facilities (Table 2).

Between January 2020 and March 2020, there was a decrease in admissions for seven of the eight conditions (Table 3). Admissions were lower for acute appendicitis (−17.4%), ACS (−57.9%), other cardiovascular disease diagnoses (−48.4%), stroke (−30%), bone fractures (−8.7%), cancer (−12.9%), and live births (−8.8%), whereas admissions were higher for respiratory tract infections (+15.1%). Respiratory tract infection admissions included 88 patients with a diagnosis of ‘coronavirus infection’ who were diagnosed upon routine testing after admission. Between March 2019 and March 2020, there was also a decrease in admissions for the same seven of the eight conditions (Table 3). Admissions were lower for acute appendicitis (−22.8%), ACS (−50.0%), other cardiovascular disease diagnoses (−81.3%), stroke (−31.3%), bone fractures (−38.8%), cancer (−23.9%), and live births (−17.0%), whereas admissions were higher for respiratory tract infections (+56.1%) (Table 3).

Unadjusted mortality rate (number of deaths recorded in a given month divided by the population in the same month) per 100,000 population was 7.91 for March 2020, 6.92 for January 2020, and 6.48 for March 2019. Total number of deaths recorded were 221 for March 2020, 192 for January 2010, and 179 for March 2019. The difference in number of deaths was not statistically significant for

Table 1
Change in number of total admissions, surgeries, daily admissions, length of stay, and in-hospital mortality (T0 = March 2020; T1 = January 2020; T2 = March 2019).

	Hamad General Hospital	Heart Hospital	National Center for Cancer Care and Research	Women's Wellness and Research Center
Total admissions				
COVID-19 period (T0)	2242	193	94	3020
Pre-COVID-19 period (T1)	2869	308	137	3465
Same period previous year (T2)	2434	293	165	3669
% Change T0 − T1	−28.0%	−59.6%	−45.7%	−14.7%
% Change T0 − T2	−8.6%	−51.8%	−75.5%	−21.5%
Elective surgeries				
COVID-19 period (T0)	312	98	10	237
Pre-COVID-19 period (T1)	526	157	11	253
Same period previous year (T2)	473	151	9	288
% Change T0 − T1	−68.6%[a]	−60.2%[b]	−10.0%[b]	−6.8%[b]
% Change T0 − T2	−51.6%[a]	−54.1%[b]	10.0%[b]	−21.5%[b]
Non-elective surgeries				
COVID-19 period (T0)	548	59	1	391
Pre-COVID-19 period (T1)	626	81	3	392
Same period previous year (T2)	661	68	1	442
% Change T0 − T1	−14.2%[a]	−37.3%[b]	−200.0%[b]	−0.3%[b]
% Change T0 − T2	−20.6%[a]	−15.3%[b]	0.0%[b]	−13.0%[b]
Average daily admissions				
COVID-19 period (T0)	72.32	6.23	3.03	97.42
Pre-COVID-19 period (T1)	92.55	9.94	4.42	111.77
Same period previous year (T2)	78.52	9.45	5.32	118.35
Change T0 − T1	−20.23	−3.71	−1.39	−14.35
Change T0 − T2	−6.2	−3.22	−2.29	−20.93
Average length of stay				
COVID-19 period (T0)	3.97	6.84	5.65	2.24
Pre-COVID-19 period (T1)	5.81	9.97	13.44	2.93
Same period previous year (T2)	6.68	12.64	13.38	3.13
Change T0 − T1	−1.84[a]	−3.13[a]	−7.79[a]	−0.69[a]
Change T0 − T2	−2.71[a]	−5.8[a]	−7.73[a]	−0.89[a]
In-hospital mortality				
COVID-19 period (T0)	1.16%	1.04%	4.26%	0.13%
Pre-COVID-19 period (T1)	0.98%	2.27%	8.76%	0.20%
Same period previous year (T2)	1.52%	2.39%	8.48%	0.14%
Change T0 − T1	0.18%[b]	−1.23%[b]	−4.50%[b]	−0.07%[b]
Change T0 − T2	−0.36%[b]	−1.35%[b]	−4.22%[b]	−0.01%[b]

[a] $P \leq 0.01$.
[b] $P > 0.05$.

comparison between March 2020 vs. January 2020 ($P = 0.2$), but was significant for the comparison between March 2020 vs. March 2019 ($P = 0.05$) (Table 4).

Discussion

To our knowledge, this is the first study to quantify the change in admission volumes and reasons for admission for common acute care conditions in the COVID-19 vs. pre-COVID-19 era. We observed a sharp decline in all studied conditions except respiratory tract infections, where an increase in admissions was observed.

Other countries have also reported a decline in hospital admission rates for various conditions. A decline in admissions for ACS was noted early in the epidemic in Italy, where the number of daily admissions for ACS to 15 hospitals declined from 18.0 in the pre-COVID-19 period to 13.3 in the COVID-19 period.[6] In the United States, overall admissions to the Department of Veterans Affairs hospitals declined by 42% in the COVID-19 period, with a significant decline noted in admissions for acute myocardial infarction, heart failure, chronic obstructive pulmonary disease, and acute appendicitis.[10] Many countries and health-care systems around the world enacted policies to reduce acute care hospital admissions by postponing elective admissions. What may not have been anticipated was a decline in non-elective admissions, which would otherwise be considered essential for optimal outcomes for the patients. Although it is possible that some patients with urgent or emergent acute conditions (e.g. acute myocardial infarction, acute appendicitis, stroke, and so on) may survive without being admitted to a hospital, lack of supervised medical care would certainly lead to overall poorer short- and long-term outcomes and a higher mortality risk. We could not identify any factor other than COVID-19 pandemic related factors to account for the decline in acute care admissions. There are several possible reasons for decline in both elective and non-elective care admissions. These include patients adhering to the physical distancing recommendations, worries about contracting the infection in a health-care facility, inability to find transportation to the hospital, critical illness which may have affected the cognitive or physical ability to seek care, and avoiding hospital visits for problems perceived to be less than critical. Another possible reason may be overly liberal criteria for admitting patients in previous months, some of which may not have been necessary. Expanding primary health-care services in Qatar may also have prevented some soft cases from being referred for admission to an acute care facility. Further studies are needed to determine the precise reasons for this decline in order to ensure that appropriate and timely care is provided to those who need such care.

Consequences of a decline in acute care admissions can be devastating for those with most severe illnesses requiring supervised medical and surgical care. We observed a reduction in admissions for multiple acute and potentially critical conditions. If patients with severe or critical illness present less frequently to acute care facilities, it is conceivable that the overall case mix index for admitted patients would be lower and reflect in terms of shorter

Table 2
Utilization metrics at four hospitals during the study periods.

	March 2020 (T0)	January 2020 (T1)	March 2019 (T2)
Hamad General Hospital			
Number of total admissions, N	2242	2869	2434
Admission and flow metrics			
Self-presentation to ED	71.10%	71.91%	70.05%
Direct admission from clinic	20.61%	18.51%	21.94%
Transfer from another hospital	0.13%	0.07%	0.16%
Other including via ambulance	8.16%	9.52%	7.85%
Disposition			
Home	92.60%	94.11%	93.14%
Another facility	2.85%	0.91%	0.74%
Died	1.16%	0.98%	1.52%
Other	3.39%	4.01%	4.60%
Heart Hospital			
Number of total admissions, N	193	308	293
Admission and flow metrics			
Self-presentation to ED	63.21%	58.44%	61.43%
Direct admission from clinic	12.95%	14.94%	11.60%
Transfer from another hospital	3.63%	5.52%	3.41%
Other including via ambulance	20.21%	21.10%	23.55%
Disposition			
Home	96.89%	94.48%	94.20%
Another facility	0%	0%	0%
Died	1.04%	2.27%	2.39%
Other	2.07%	3.25%	3.41%
National Center for Cancer Care and Research			
Number of total admissions, N	94	137	165
Admission and flow metrics			
Self-presentation to ED	45.74%	46.72%	39.39%
Direct admission from clinic	40.43%	32.85%	37.58%
Transfer from another hospital	0.00%	0.73%	0.61%
Other including via ambulance	13.83%	19.71%	22.42%
Disposition			
Home	92.55%	86.13%	89.09%
Another facility	1.06%	1.46%	0.61%
Died	4.26%	8.76%	8.48%
Other	2.13%	3.65%	1.82%
Women's Wellness and Research Center			
Number of total admissions, N	3020	3465	3669
Admission and flow metrics			
Self-presentation to ED	39.37%	40.61%	40.45%
Direct admission from clinic	16.52%	14.66%	17.23%
Transfer from another hospital	0.07%	0.17%	0.14%
Other including via ambulance	44.04%	44.56%	42.19%
Disposition			
Home	99.21%	98.96%	99.24%
Another facility	0.26%	0.40%	0.22%
Died	0.13%	0.2%	0.14%
Other	0.40%	0.43%	0.41%

ED, emergency department.

length of stay and lower in-hospital mortality. We did indeed observe a significantly shorter length of stay across all hospitals included in the study. While in-hospital mortality trend was observed to be numerically lower for most comparisons, none of it reached statistical significance likely due to small number of in-hospital deaths. A concomitant trend toward a higher overall unadjusted death rate in the country was observed, although this must be interpreted with extreme cautions because we only

Table 3
Change in number of patients with selected diagnoses across four hospitals during the study periods.

	March 2020 (T0)	January 2020 (T1)	March 2019 (T2)	T0 − T1%	T0 − T2%
Acute appendicitis	92	108	113	−17.4%	−22.8%
Acute coronary syndrome	114	180	171	−57.9%	−50.0%
Other cardiovascular disease diagnoses	64	95	116	−48.4%	−81.3%
Stroke	83	108	109	−30.1%	−31.3%
Bone fracture(s)	183	199	254	−8.7%	−38.8%
Cancer	155	175	192	−12.9%	−23.9%
Live births	1291	1405	1510	−8.8%	−17.0%
Respiratory tract infections[a]	337[b]	286	148	15.1%	56.1%

[a] excluding tuberculosis.
[b] includes 88 cases of coronavirus disease.

Table 4
Population statistics for the State of Qatar for the study periods.

	March 2020 (T0)	January 2020 (T1)	March 2019 (T2)	T0 - T1%	T0 - T2%
Total population	2,795,484	2,773,221	2,760,586	0.8	1.2
Total deaths	221	192	179	13.1[a]	19.0[b]
Unadjusted death rate	7.91	6.92	6.48	12.5	18.1

[a] $P = 0.2$.
[b] $P = 0.05$.

studied three months data and did not ascertain the long-term trends or the reasons for this variation. This is an important hypothesis-generating observation and must not be interpreted as a causal link. The overall crude mortality rate in the country is particularly low reflecting the demographic pattern and the young population of Qatar.

The increase in admissions due to respiratory tract infections is likely a reflection of heightened awareness of COVID-19 infection and increased vigilance practiced by health-care practitioners. We did observe an increase in number of patients who were admitted under isolation precautions in the March 2020, which supports this impression (data not shown). The four hospitals included in this study were designated to be COVID-19-free facilities, and all diagnosed patients were admitted or transferred to designated facilities for COVID-19 patients. The number of admissions due to respiratory tract infections may have been higher if COVID-19 patients had free access to these hospitals. This policy likely kept the nosocomial COVID-19 infection rate and infection among health-care workers at a very low level in these facilities. Cohorting of patients also led to more efficient use of resources, both in terms of health-care personnel and equipment.

Our study raises important questions about future pandemic planning. Follow-up studies are needed to determine whether this observed decline in admissions will translate to more severe presentations, late diagnosis, or long-term disability related to these diagnoses. For example, determining the proportion of patients presenting with ruptured appendix, more advanced cancer, more residual weakness or disability after stroke, and higher rates of advanced heart failure can lead to policies targeting such patients in any future pandemic setting, and devising strategies to get them in appropriate care setting earlier. It is also important to understand the magnitude and burden of these consequences when current travel and physical distancing restrictions are removed. Knowing this burden will be critical in planning for the possible surge of patients once these restrictions are lifted.

The strengths of our study include evaluating multiple hospitals which see a variety of acute care conditions and are national referral centers. A variety of acute care conditions were studied to provide a broad understanding of the change in admission patterns. We also studied national mortality trends in an attempt to understand large scale consequences of our findings. Despite numerous strengths, these data need to be interpreted with caution, as they represent only a snapshot in time. Particularly, any correlation with national mortality statistics must be interpreted with extreme caution because numerous factors may affect those statistics. No causal inference can be drawn from the present study regarding the impact of our findings on overall mortality rates for the entire country.

In summary, we observed a significant decline in hospital admissions across several hospitals in a national health-care system, with a significant decline in admissions due to seven of eight acute care conditions studied. This decline was associated with a shorter length of stay but not associated with a change in in-hospital mortality rate. A possible small increase in unadjusted mortality rate in the country requires further study to determine if there is any correlation with the change in hospital admission rates.

Author statements

Ethical approval

The study was reviewed and approved by the Institutional Review Board at Hamad Medical Corporation (MRC-05-034).

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Competing interests

None declared.

Author contributions

A.A.B. reports contributing to concept and study design; data acquisition. A.A.B., A.B.K., and A.A. reports contributing to the drafting of the manuscript. A.A.B. and A.A. reports contributing to data analysis. A.A.B., A.B.K., N.A.M., A.M.A., N.A.A., M.U.A.H., H.A.H.S., R.B., and A.A. reports contributing to data interpretation and critical appraisal and review.

Authorship statement

Dr. Butt had complete access to data at all times and accepts the responsibility of the integrity of this article.

Disclaimer

The views expressed in this article are those of the authors and do not necessarily represent official government views or policy of the State of Qatar or Hamad Medical Corporation.

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Original Research

Multiple deprivation and geographic distance to community physical activity events — achieving equitable access to parkrun in England

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ABSTRACT

Objectives: To evaluate geographic access to free weekly outdoor physical activity events ('parkrun') in England, with a particular focus on deprived communities, and to identify optimal locations for future events to further maximise access.
Study design: This study is a cross-sectional ecological analysis of the socio-economic disparities in geographic access to parkrun events in England in late 2018.
Methods: We combined geolocation data on all English Lower Layer Super Output Areas and parkrun events to calculate geodesic distances to the nearest event for more than 32,000 communities in England. We use this measure of geographic access to summarise the relationship between access and socio-economic deprivation, measured using the index of multiple deprivation. We then used geographic coordinates of public green spaces in England to conduct a simple location-allocation analysis to identify 200 locations for future event locations that would maximise access.
Results: In England, 69% of the population live within 5 km of one of the 465 parkrun events. There is a small negative correlation between distance and deprivation, indicating that access is slightly better in more socio-economically deprived areas. Setting up an additional 200 events in optimal locations would improve access: the average distance to the nearest parkrun event would improve by 1.22 km, from 4.65 km to 3.43 km, and approximately 82% of the English population would live within 5 km of a parkrun event.
Conclusion: Over two-thirds of the English population live within 5 km of a parkrun event, and contrary to our expectation, we find that geographic access is slightly better for those living in more deprived communities. Creating additional events may improve geographic access, but effective strategies will still be needed to increase engagement in new and existing events by those living in socio-economically deprived areas.

Introduction

Insufficient physical activity is one of the leading causes of disease and disability worldwide.[1] In the UK, around one in six deaths is attributable to low levels of physical activity.[2] It is also a major contributor to health inequalities, as people from low socio-economic backgrounds are both disproportionately likely to be inactive[3,4] and be affected by physical inactivity-related diseases.[5] Increasing the physical activity levels of the population is therefore high on the public health agenda: it not only has the potential to improve quality of life, reduce mortality rates and alleviate the strain on health and social care services but also reduce the gap in health inequalities.[6]

However, designing effective public health interventions that increase population physical activity is a considerable challenge.[7,8] Implementing such interventions in a way that does not increase health inequalities might even be more difficult. Studies have shown that programmes to increase physical activity often fail to reach deprived communities and those most in need, suggesting

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access to physical activity facilities shows spatial and social inequality.[9,10] In this regard, parkrun, an international movement which organises free weekly 5 km running and walking events in public spaces, might provide valuable lessons.

Since its founding in 2004, as a small event in London with 13 participants, parkrun has grown to become one of the world's largest mass sporting events, with up to 360,000 participants in more than 20 countries.[11,12] The volunteer-led events are often characterised as accessible and inclusive.[13,14] The organisation has been widely praised as being successful in encouraging participation particularly in individuals who were previously inactive.[15,16]

Notwithstanding these subjective accounts, the expansion of parkrun in England, as elsewhere, has been largely grassroots, driven by *demand* rather than *need*. It might therefore be the case that parkrun events are primarily located in areas that are less deprived, while people living in more deprived communities may not have the same opportunities to participate. In 2019, Sport England announced funding to support the creation of 200 new parkrun events across England within three years, with the specific aim of increasing participation of individuals from lower socio-economic groups.[17]

The aims of this study are two-fold: first, to evaluate whether geographic access to parkrun events in England is equitable across areas with different levels of deprivation; and second, to identify 200 optimal locations for future events to improve geographic access, in particular for deprived communities.

Methods

Study design

This study is a cross-sectional ecological analysis of the socio-economic disparities in geographic access to parkrun events in England at the end of 2018. All analyses were conducted on the level of Lower Layer Super Output Areas (LSOAs), which divide England into 32,844 geographic units which, on average, have a population of approximately 1700. We assessed the relationship between access, defined as the distance (as the crow flies) to the nearest parkrun event, and socio-economic deprivation, measured using the Index of Multiple Deprivation (IMD). In addition, we used information on public green spaces in England to conduct a simple location-allocation analysis, to identify 200 locations for future parkrun events that maximise access for the population.

Data sources

For this study, we combined data on three types of geospatial entities: (1) LSOAs, (2) parkrun events, and (3) public green spaces.

1) For all 32,844 LSOAs, we retrieved geographic locations, defined by the coordinates of its population-weighted centroid; 2017 total population estimates; and the 2015 IMD from the Office for National Statistics.[18–20]
2) We included all 465 public parkrun events which were in operation in England by December 12th, 2018 – on this date, Sport England announced their plan to provide funding for setting up 200 additional parkrun events across England.[21] The locations of the events were obtained from the parkrun UK website.[22]
3) The locations of public green spaces in England were retrieved from an open data set of Ordnance Survey.[23] Parkrun events are held in various settings and terrains and do not always require a single 5 km loop – some events have courses that involve running a combination of shorter loops. After evaluating existing parkrun event courses, we decided to consider all public parks, gardens and playing fields in England with an area of 0.1 km^2 or more potentially suitable for hosting events (n = 2842).

Variables

The two variables of interest were access to parkrun events and deprivation of LSOAs.

Access to parkrun was defined as the geodesic distance (as the crow flies) from LSOA's population-weighted centroid to its nearest event. For each of the 32,844 LSOAs, we computed the geodesic distances between its population-weighted centroid and all 465 parkrun events that were in operation on December 12th, 2018 and then selected the shortest distance.

The socio-economic deprivation of LSOAs was measured using the 2015 IMD. It is a measure of relative deprivation, which has been used in many similar studies. The IMD combines 37 indicators from seven domains (income, employment, education and skills, health and disability, crime, housing and services, and living environment) into a single score. The score ranges from 0 (least deprived) to 100 (most deprived).[24]

Other covariates, which are likely to affect the availability of parkrun events within an area (e.g. population density or demographics), were not taken into account because we did not aim to assess to what extent deprivation independently 'explained' access. Rather, we sought to evaluate whether or not people living in deprived areas have better or worse geographic access, under the actual circumstances.

Analysis

Mean, standard deviation, median, interquartile range, and range were used as descriptive statistics. We then assessed the association between the IMD and the distance to the nearest parkrun event on the LSOA level. Our hypothesis was that more socio-economically deprived areas had worse access, i.e. longer geodesic distances to the nearest parkrun event than less deprived areas. Pearson and Spearman correlation coefficients were computed using the LSOAs' total population as weights. We also conducted a stratified analysis, for which we grouped LSOAs into IMD quintiles (most, more, median, less, least deprived) and assessed access to parkrun events in each stratum.

Identifying optimal locations for new parkrun events

We conducted a location-allocation analysis to solve the following problem. Parkrun UK received funding to start 200 additional parkrun events; there are 2842 public green spaces in England in which new events could be set up – which 200 locations should be selected, to maximise access for the greatest number of people?

More specifically, the objective was to minimise the population-weighted total sum of distances between all LSOAs and their nearest parkrun event. To identify the optimal 200 green spaces, we applied a simple greedy algorithm that consisted of two steps. Firstly, for each green space, we evaluate how setting up a parkrun event would affect the sum of distances, given the locations of all existing events (i.e. for how many LSOAs this green space would be the nearest parkrun event, and by how much it would decrease the respective distances). Second, the green space with the greatest effect is selected and added to the set of existing parkrun events. This procedure is repeated 200 times.

More formally, the first step of the algorithm evaluates the following expression:

$$\underset{c \in C}{argmin} \sum_{i=1}^{32,844} d_i(E \cap c)*p_i$$

The function yields the candidate green space c, from the set of all 2842 green spaces C, which minimises the sum of the population-weighted distances between LSOAs and their nearest parkrun event. The total population of LSOA i is denoted p_i, and $d_i(E \cap c)$ denotes LSOA i's distance to the nearest parkrun event, which can either be an existing event from the set of 465 parkrun events, denoted E, or the candidate green space c, whichever is nearest.

To identify the optimal new locations for setting up 200 new parkrun events consecutively, the selection procedure is repeated 200 times. At each step, the single best candidate green space location is selected, added to the set of established parkrun events E and removed from the set of available green spaces C. This means, the effect of the green space selected at step k is taken into account when selecting the kth+1 location.

We assessed the overall impact of setting up 200 new parkrun events on the geographic access to parkrun events in England. We also investigated the effects on LSOAs across IMD quintiles in a distributional analysis.

Data and source code availability

All data and the R source code that were used to generate the results of this study are provided on an open repository.[25] Ethical approval

Ethical approval was obtained from the Sheffield Hallam University Ethics Committee (ER10776545). We did not collect any personal information, and only used aggregate secondary data. The parkrun Research Board approved this research project, and four of its members (A.M.B., H.Q., E.G., SS.J.H.) were actively involved in the interpretation of findings and writing of this manuscript.

Results

Descriptive statistics

As of 12th December 2018, approximately 7%, 69%, and 91% of the English population lived within 1, 5, and 10 km distance of a parkrun event, respectively. Only 578,043 people (1% of the English population) lived more than 20 km from an event. The mean (standard deviation (SD)) and median (interquartile range (IQR)) distance to the nearest parkrun event were 4.65 (4.22) and 3.39 (1.99–5.83) km. The largest distance was observed for the 2259 people living on the Isles of Scilly, who live about 76 km away from the next parkrun event on the mainland. On average, each parkrun event is the closest event for 71 LSOAs (43), with a combined population of 119,612 (74,290). Further descriptive statistics are provided in Table 1.

Association between deprivation and access

There was a negative relationship between IMD and the distance to the nearest parkrun event: the (population-weighted) Pearson and Spearman correlation coefficients were −0.15 and −0.18, indicating a small negative correlation. This means that more deprived LSOAs tended to have shorter distances to the nearest parkrun event, i.e. better geographic access.

The analysis of distances by IMD quintile in Table 2 shows that people living in the 20% most deprived LSOAs had the best geographic access, with a mean and median distance to the nearest parkrun event of 3.51 and 2.79 km, respectively. Depending on the metric, the worst access was observed for LSOAs in the middle (mean distance = 3.36 km) or the less deprived group (median distance = 3.93 km). Further results of the distributional analysis are provided in Table 2.

Optimal locations for new parkrun events

Fig. 1 shows the parkrun events (blue circles) that existed on 12th December 2018 alongside recommendations for 200 additional event locations (red triangles), which minimise the sum of the population-weighted geodesic distances from the LSOA centroids, i.e. maximise overall access to parkrun for the greatest number of people. The numbers correspond to the rank, where 1 is the location which would improve access the most. The names and exact locations of the selected 200 green spaces are provided in Table S1 in the appendix. We also created an interactive map, which can be accessed online, to explore the locations of existing and recommended parkrun event locations in more detail: https://bitowaqr.github.io/parkrun_access_equity/.

We estimated that setting up new parkrun events in those 200 green spaces would improve access for around 16.5 million people (30% of the population) from 9854 LSOAs. For these people, the distance to the nearest event would, on average, be reduced by 4.09 km (SD = 3.97). Overall, it would reduce the average and median distance to the nearest parkrun event from 4.65 and 3.39 km to 3.43 and 2.59 km. The percentage of people who live within 5 km of a parkrun would increase from 69% to 82%.

The distributional analysis in Table 2 shows, for each IMD quintile, geographic access under the current situation (12th December 2018) and after the creation of 200 new events. Overall, setting up 200 new events in the recommended green spaces would amplify the negative socio-economic gradient in geographic access. The population-weighted Pearson and Spearman correlation coefficients changed from −0.15 and −0.18 before, to −0.20 and −0.23 afterwards, indicating that improvements in access to parkrun events were greater for more deprived LSOAs than less deprived LSOAs. Nevertheless, the distributional analysis showed that the improvement in access was smallest for LSOAs in the most deprived quintile.

Table 1
Descriptive statistics of LSOAs and parkrun events.

Variable	Mean (SD)	Median (Q25-Q75)	Range
LSOAs (n = 32,844)			
Population	1693 (405)	1612 (1452; 1834)	362-13,404
IMD	21.67 (15.59)	17.40 (9.65; 30.07)	0.48–92.60
Distance (in km) to the nearest event	4.65 (4.17)	3.39 (1.99; 5.83)	0.04–76.44
Parkrun events (n = 465)			
Catchment area[a] population	119,612 (74,290)	103,952 (68,837; 151,488)	7855–628,010
Catchment area[a] LSOAs	71 (43)	62 (40; 87)	6–350

[a] Number of LSOAs/total population for which a given parkrun event is the nearest.

Table 2
Distributional analysis. The table shows the distance (in km) to parkrun events before and after 200 new parkrun events are set up at optimal green spaces, stratified by IMD quintiles.

	Current situation (December 12th, 2018)			After 200 new parkrun events are set up		
Variable	Mean (SD)	Median (Q25-Q75)	Range	Mean (SD)	Median (Q25-Q75)	Range
Least deprived	4.93 (3.62)	3.91 (2.27; 6.67)	0.12–58.54	3.79 (2.61)	3.09 (1.92; 4.96)	0.12–25.58
Less deprived	5.21 (4.24)	3.93 (2.28; 6.99)	0.14–76.44	3.92 (3.04)	2.99 (1.84; 5.09)	0.14–48.02
Median deprived	5.36 (5.01)	3.68 (2.12; 6.83)	0.11–60.81	3.98 (3.55)	2.79 (1.70; 4.91)	0.11–33.74
More deprived	4.26 (4.38)	2.96 (1.76; 5.00)	0.04–59.44	3.03 (2.78)	2.27 (1.47; 3.49)	0.04–24.07
Most deprived	3.51 (3.01)	2.79 (1.71; 4.39)	0.07–36.17	2.43 (1.68)	2.12 (1.41; 3.02)	0.05–24.30
Overall	4.65 (4.17)	3.39 (1.99; 5.83)	0.04–76.44	3.43 (2.86)	2.59 (1.63; 4.16)	0.04–48.02

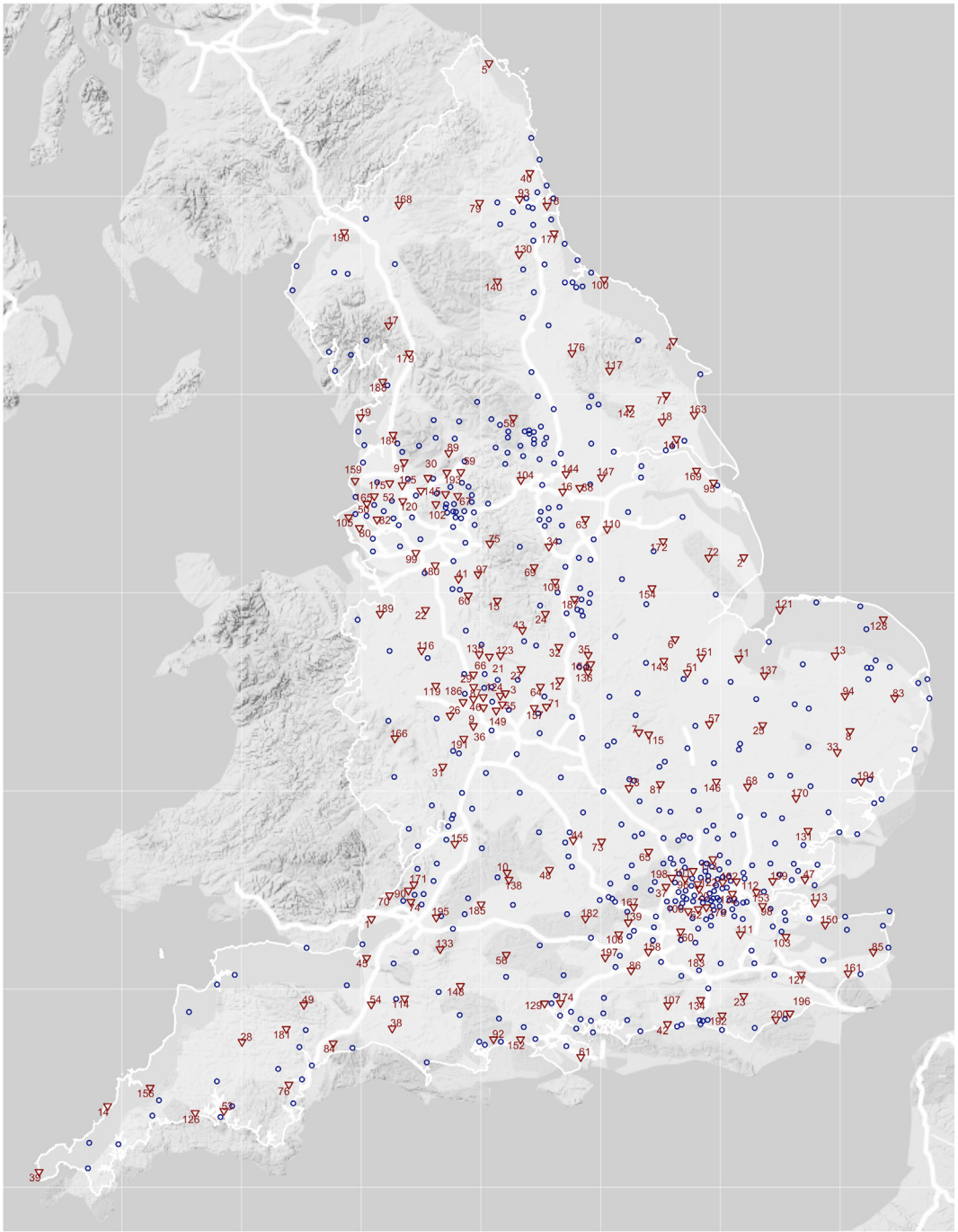

Fig. 1. Map of England showing current parkrun events (blue circles) and recommended new event locations (red triangles) ranked in descending order of estimated effect on overall population-weighted access. Information on all 200 identified optimal green space locations are provided in the appendix. (For interpretation of the references to colour in this figure legend, the reader is referred to the Web version of this article.)

Discussion

As of 12 December 2018, the median distance to the nearest parkrun event was 3.39 km and more than two-thirds of the English population lived within 5 km (the parkrun distance) of a parkrun event. Contrary to our expectation, we did not find that access was better for people living in less deprived areas. In fact, those living in the most deprived areas had the best geographic access to parkrun – it is rare in public health for inequalities to exist in this direction.[10,26]

Our analysis has shown that setting up 200 new events in the recommended (optimal) green spaces in England would reduce the average distance to the nearest parkrun event by 1.22 km, increasing the percentage of English residents who live within 5 km of a parkrun to 82%. Moreover, the recommended expansion of parkrun would improve the geographic access for the most deprived areas more than the access of those living in more affluent areas.

The main finding, that geographic access to parkrun events is better in more deprived communities, is surprising. Parkrun events are set up by volunteers, based on demand not need. Studies have shown that the level of physical activity, and the availability of physical activity facilities generally declines with the level of deprivation.[4] Opportunities for physical activity are often lacking in areas in most need.[27] Parkrun events, in contrast, seem to be often held in or near deprived areas, and are free to attend, giving anyone equal access, irrespective of their socio-economic background. Nevertheless, in a previous analysis, we found that participation in parkrun has a strong socio-economic gradient with considerably higher participation rates in less deprived areas: about a third of all participants came from LSOAs in the least deprived quintile, whereas only 7% came from the most deprived quintile.[28,29] This suggests that providing the opportunity to participate in parkrun events, while a necessary first step to enable participation, has not been sufficient to engage people living in deprived communities.[9] This means, creating additional events in optimal locations could improve overall geographic access further, yet effective strategies will still be needed to increase equity in engagement in new and current events.

There are several strengths of this study that deserve mention. First, it is the first study of geographic access to parkrun in England – therefore the approach is novel and the data untapped. Second, the analysis makes use of large and rich data sets, with more than 30,000 LSOA and more than 400 existing parkrun events; it is unlikely that individual outliers are affecting the results. The almost universal availability of parkrun events throughout the country provides a learning opportunity to explore socio-spatial determinants influencing physical activity behaviour on a national scale.[16] Our study contributes to the limited research in this area and identifies possible leads for further investigation.

However, there are also limitations. Most importantly, geographic access is not measured as travel distance, or travel time, but as geodesic distance. In some cases, for example, where natural barriers such as hills, lakes or rivers block routes, the actual distance travelled may be far in excess of the geodesic distance.[30]

Furthermore, the list of green spaces that we considered as potential sites for future parkrun events is neither comprehensive nor without limitations. Not all included green spaces may be suitable to host events (e.g. because of the terrain or the setting), and the list also does not contain all suitable places (e.g. many blue spaces such as beaches and promenades are not included).

Finally, our analysis has been concerned only with determining to what extent deprived communities have geographic access to parkrun events. We did not investigate what other factors independently explain access more generally. It should be noted, however, that a contributing factor for the negative relationship between IMD and access is likely to be population density: deprived areas cluster in urban areas, where also most parkrun events take place. Rural areas, on the other hand, may therefore have worse geographic access. Further studies are required to better understand wider determinants of access to parkrun and/or physical activity facilities more generally.

Studying barriers to participation in parkrun, other than geographic access, is likely to improve our understanding of the reasons why physical activity levels are lower in more deprived areas and may help to design more effective public health interventions to increase levels of physical activity in the population. Future research should build on this work and develop a model to assess the (cost-)effectiveness of setting up new events, and other strategies, not only in terms of improved potential access but also actual participation. This requires estimating the causal and marginal effects of different interventions on participation, and therefore physical activity levels, using longitudinal data and sophisticated modelling techniques.

Conclusion

In England in December 2018, 69% of the population lived within 5 km of a parkrun event. Creating 200 new events in the recommended (optimal) green spaces would further improve access, increasing this to 82%. Contrary to our expectation, we find that geographic access is slightly better for those living in more deprived communities. Given that participation rates are generally lower in deprived areas, improving access alone seems unlikely to significantly reduce inequalities in participation and physical activity. To design more effective strategies to improve engagement from deprived communities, a deeper understanding of the barriers to taking part in mass participation physical activity events is needed.

Author statements

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Ethical approval

Ethical approval was obtained from the Sheffield Hallam University Ethics Committee (ER10776545). The authors did not collect any personal information but only used aggregate secondary data. The parkrun Research Board approved this research project, and four of its members (A.M.B., H.Q., E.G. and S.J.H.) were actively involved in the interpretation of findings and writing of this manuscript.

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Competing interests

A.M.B., H.Q., E.G. and S.J.H. are members of the parkrun Research Board.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2020.09.002.

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Short Communication

Non–COVID-19 excess deaths by age and gender in the United States during the first three months of the COVID-19 pandemic

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ABSTRACT

Objectives: The first three months of the COVID-19 pandemic has disrupted healthcare systems, creating an environment by which deaths have occurred that are not directly due to COVID-19, but have occurred owing to the healthcare and societal environment resulting from COVID-19. The objective of this research is to quantify such excess deaths, partitioned by age group and gender.
Study design: This is a data analysis.
Methods: Excess deaths by age and gender are estimated using provisional death data available from the Centers for disease control and prevention (CDC) over the time period from March 1, 2020 through May 30, 2020. Previous year fatality and population data are used as the benchmark.
Results: Several of the eighteen age and gender cohorts experienced statistically significant excess deaths. The results also indicate that COVID-19 has been protective for one of the age and gender cohorts.
Conclusions: There have been more excess deaths in several age group and gender cohorts during the first three months of the pandemic, beyond direct deaths directly attributable to COVID-19. These non–COVID-19 excess deaths are most apparent in the 25- to 44-year age group for women and 15- to 54-year age group for men. Further research is needed to assess the cause of such excess deaths and introduce safeguards to reduce such deaths in the future.

Introduction

The COVID-19 pandemic has resulted in more than 7 million infections and more than 210,000 deaths in the US. In mid-March, state governments began to close their economies and encouraged citizens to stay home to suppress the spread of the virus. The purpose of such shelter-in-place orders was to better facilitate physical distancing to reduce community virus transmission, blunt the anticipated surge in demand of hospital intensive care unit beds and ventilators and gain time to provide sufficient personal protective equipment for healthcare workers. The preponderance of cases and deaths from mid-March 2020 through mid-June 2020 was in large, densely populated urban areas such as New York City, Chicago and Detroit. This surge mostly sidestepped rural communities during this period, creating an uneven impact of COVID-19 across the nation.

Given the speed at which COVID-19 spread in early March 2020 and the uncertainty of its virulence, the aggressive step of closing the US economy was prudent and widely accepted. Hsiang et al.[1] estimated that without such actions, 4.8 million additional confirmed cases would have occurred in the US alone through May 2020. Using a 5% case fatality rate, this would have translated into 240,000 additional deaths.

Through mid-June, data collected and disseminated by the CDC have demonstrated a clearer picture of which population cohorts are most vulnerable to COVID-19 (such as those older than 65 years and those with underlying health conditions). This has provided a road map for protecting at-risk people while progressing towards reopening communities and local economies. To limit the spread of the virus without shelter-in-place orders, in addition to testing and contact tracing, public health countermeasures of hand hygiene, physical distancing and face coverings provide the best available defences to limit virus transmission and protect the most vulnerable populations.

The CDC disseminates a weekly summary of provisional deaths from all causes and COVID-19 deaths, broken down by age and

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Table 1
Average weekly death statistics (female).

Age cohort (years)	2020 total deaths	2020 COVID-19 deaths	2020 non–COVID-19 deaths	Hybrid 2019 deaths	Pooled standard deviation estimator	*P*-value
5–14	37.5	0.2	37.3	45.3	1.7	<0.001*
15–24	166.8	3.8	163.0	155.2	4.4	0.05
25–34	412.5	17.6	394.9	346.7	8.5	<0.001
35–44	673.2	42.4	630.8	561.1	8.8	<0.001
45–54	1345.5	122.0	1223.5	1200.7	22.2	0.16
55–64	3245.2	340.7	2904.5	2823.9	66.9	0.13
65–74	5420.4	662.0	4758.4	4586.1	109.0	0.07
75–84	7706.3	1005.9	6700.4	6506.7	199.0	0.17
85+	12,568.2	1688.5	10,879.7	10,436.8	425.5	0.16

gender.[2] The disruptive social and economic upheavals created by the COVID-19 pandemic have led to excess deaths that are either directly or indirectly attributable to COVID-19. The CDC reports estimates of such excess deaths.[3] This article uses CDC estimates for 2019 deaths as a baseline to estimate excess deaths specified by age and gender cohorts. This analysis provides an alternative perspective by which to estimate excess deaths and the health impact of COVID-19.

Methods

Provisional death data reported by the CDC from March 1, 2020 through May 30, 2020, are used to estimate the mean and standard error of the number of weekly deaths, both from all causes and those attributed to COVID-19 across 18 cohorts, broken down by age and gender (male and female).[2] Label the data from this 13-week period as '2020 weekly deaths'.

Point and standard error estimates for the expected number of deaths per week for each age and gender cohort are computed based on CDC data for 2018 death rates[4] and 2019 United States Census Bureau population estimates,[5] labelled 'hybrid 2019 weekly deaths'. This represents the most recent age-based mortality and population data available, hence can serve as a benchmark for assessing 2020 excess deaths. The standard errors were estimated using the 2018 monthly deaths for each age and gender,[4] the most recent death data available from the CDC.

For each age and gender cohort, a one-sided Student t-test was used to test the null hypothesis that the expected 2020 non–COVID-19 weekly deaths are equal to the expected hybrid 2019 weekly deaths, compared with the alternative hypothesis that it is greater. The 2020 non–COVID-19 weekly death estimates are computed by subtracting 2020 COVID-19 weekly deaths from all 2020 weekly deaths. A pooled standard error estimator for the Student t-test was computed by taking the square root of the sum of the standard error squared for the 2020 non–COVID-19 weekly deaths plus the standard error squared for the 2018 weekly deaths (rescaled used the 2018 monthly deaths).

Results

Table 1 shows estimates for the expected 2020 weekly female deaths, 2020 COVID-19 weekly female deaths, 2020 non–COVID-19 female deaths, 2019 hybrid weekly female deaths, pooled standard error estimator for weekly female deaths and *P*-values for the Student t-test statistic. The data in Table 2 are depicted in a similar manner for males.

The *P*-values create a hierarchy for evaluating excess deaths. An age cohort is labelled 'statistically significant' if the *P*-value is lower than 0.001, indicating strong evidence that the expected 2020 non–COVID-19 weekly deaths are larger than the expected hybrid 2019 weekly deaths. A cohort is labelled 'statistically inconclusive' if the *P*-value is between 0.001 and 0.05, indicating marginal evidence that the expected 2020 non–COVID-19 weekly deaths are more than the expected hybrid 2019 weekly deaths. A cohort is labelled 'statistically insignificant' if the *P*-value is higher than 0.05, indicating weak evidence that the expected 2020 non–COVID-19 weekly deaths are more than the expected hybrid 2019 weekly deaths.

For 17 of the 18 age and gender cohorts, the 2020 non–COVID-19 average weekly deaths are more than the hybrid 2019 average weekly deaths. The one exception is the 5- to 14-year age group for females, which indicates that the 2020 non–COVID-19 average weekly deaths are fewer; hence, COVID-19 was protective for these young girls (*P*-value < 0.001 labelled with a '*' in Table 1).

For women, two age cohorts (25–34 and 35–44 years) show a statistically significant (*P*-value < 0.001) increase in expected 2020 non–COVID-19 weekly deaths compared with the hybrid 2019 weekly deaths. For men, four age cohorts (15–24, 25–34, 35–44 and 45–54 years) show a statistically significant increase in expected 2020 non–COVID-19 weekly deaths. Alternatively, for males aged 5–14, 75–84 and ≥85 years and for females aged 15–24,

Table 2
Average weekly death statistics (male).

Age cohort (years)	2020 total deaths	2020 COVID-19 deaths	2020 non–COVID-19 deaths	Hybrid 2019 deaths	Pooled standard deviation estimator	*P*-value
5–14	60.5	0.9	59.5	59.0	2.6	0.41
15–24	477.1	7.4	469.7	418.3	11.7	<0.001
25–34	964.6	40.1	924.5	789.2	23.7	<0.001
35–44	1317.7	106.1	1211.6	995.2	24.4	<0.001
45–54	2335.2	288.3	2046.9	1902.5	35.8	<0.001
55–64	5303.5	672.3	4631.2	4405.7	84.8	0.01
65–74	7626.4	1095.3	6531.1	6206.0	148.9	0.02
75–84	8399.4	1220.5	7178.9	6912.6	205.0	0.11
85+	7827.3	1084.8	6742.5	6610.4	253.3	0.31

45–54, 55–64, 65–74, 75–84 and ≥85 years, the expected 2020 weekly death increases may be explained by COVID-19 because increases in non–COVID-19 deaths were statistically insignificant (*P*-value > 0.05).

Discussion

The data do not explain why there is a statistically significant increase in expected 2020 non–COVID-19 weekly deaths compared with the expected hybrid 2019 weekly deaths. Czeisler et al.[6] discuss delays or avoidance of non–COVID-19 medical care during the pandemic, which could contribute to excess deaths beyond those attributed to COVID-19.

For all but one age and gender cohort, there were more 2020 average weekly deaths than the hybrid 2019 average weekly deaths. One possible explanation for this is that 2019 weekly deaths are not uniformly distributed across the year, which is highly likely. Another explanation is that because the 2020 population has a higher is larger than 2019, there may be more deaths, although this increase is likely to be negligible compared with the actual number of deaths. To overcome these limitations, we used a *P*-value cut-off of 0.001 (rather than 0.05) to assess statistical significance and a cut-off of 0.05 to assess statistical insignificance, whereas all other values in between were classified as statistically inconclusive.

The CDC provides weekly updates of provisional death reports,[2] which continue to be adjusted for the time period from March 1, 2020, through May 30, 2020, as new data become available. Therefore, the values reported in Table 1 will continue to change, albeit slightly as at from the time of the analysis. Because only new deaths are added, this will tend to result in *P*-values getting marginally lower.

The key takeaway from this analysis is that excess deaths across multiple age and gender cohorts occurred beyond what has been attributed to COVID-19. These excess deaths indicate that people across many age and gender cohorts have died unexpectedly. Over the ensuing months, possible explanations for such excess deaths may become more apparent.

Author statements

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Ethical approval

No ethical approval was required for this study. The analysis uses only publicly available data reported in the literature.

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No funding was required for this study. The analysis uses only publicly available data reported in the literature.

Competing interests

The authors declare no conflicts of interest.

Author contributions

Both authors contributed to the ideas that led to the article. S.H.J. contributed to the statistical analysis, the literature review and the manuscript preparation. J.A.J. contributed to the concept and provided background on infectious diseases and public health. S.H.J. wrote the first draft of the article. J.A.J. provided extensive feedback and comments. Both authors read and approved the final version of the manuscript.

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Original Research

Predictors of laboratory-positive COVID-19 in children and teenagers

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ABSTRACT

Objective: The aim of the study was to identify factors predicting laboratory-positive coronavirus disease 2019 (COVID-19) in pediatric patients with acute respiratory symptoms.
Study design: We conducted a cross-sectional analysis of a prospective cohort study.
Methods: Data from 1849 individuals were analyzed. COVID-19 was confirmed (reverse transcription-quantitative polymerase chain reaction) in 15.9% of patients, and factors predicting a positive test result were evaluated through prevalence odds ratios and 95% confidence intervals.
Results: Increasing age, personal history of obesity, and household contact with a case were found to be associated, in the multiple regression model, with increased odds of a positive test result. Young patients residing in areas with higher population sizes, as well as those with severe respiratory symptoms, were less likely to be laboratory confirmed.
Conclusions: Early identification and isolation of children and teenagers with suggestive symptoms of COVID-19 is important to limit viral spread. We identified several factors predicting the laboratory test result. Our findings are relevant from a public health policy perspective, particularly after the restart of in-person academic activities.

Introduction

The coronavirus disease 2019 (COVID-19) pandemic caused by severe acute respiratory syndrome coronavirus 2 (SARS-CoV-2) has evolved quickly around the globe, and the related socio-economic burden is high.[1] Less than 2% of COVID-19 cases are reported in pediatric patients, and they commonly show milder symptoms and a better prognosis than adults.[2]

Currently, it is unclear if this low rate among children and teenagers results from a diminished susceptibility to SARS-CoV-2 infection or from a higher prevalence of asymptomatic cases.[3] Nevertheless, there is a general consensus regarding the role of young individuals in the spread of viral respiratory pathogens.[4,5] This highlights the relevance of preventive interventions including early case identification and quarantining and limitating crowded physical activities.[6]

We aimed to evaluate predictors of laboratory-positive SARS-CoV-2 infection among children and teenagers with symptoms of acute viral respiratory infection.

Methods

Study population

We conducted a cross-sectional analysis of a nationwide cohort study. Suspected COVID-19 cases (disease onset: February–August 2020) from any age were enrolled in an ambispective cohort study that served as the source for the study sample. Individuals were followed up until disease classification and clinical outcomes. Participants were identified from the nominal records of the National System of Epidemiological Surveillance of Mexico, which operates as per normative standards.[7]

Potentially eligible children (aged younger than 12 years) and teenagers (12–15 years old) were those with conclusive results (confirmed SARS-CoV-2 infection, no/yes) of reverse transcription-

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quantitative polymerase chain reaction (qRT-PCR) and symptom onset from February to April 2020. Asymptomatic patients (without fever, rhinorrhea, or cough; $n = 3$) at the collection of clinical specimens and those without complete data of interest were excluded ($n = 11$). All young individuals fulfilling the eligibility criteria were included in our analysis. Enrolled subjects sought for medical attention at any of more than 1400 healthcare settings (all three levels of care) belonging to the Mexican Institute of Social Security (IMSS, the Spanish acronym) located all across the country.

Data collection

Clinical and epidemiological data of interest were obtained from the audited surveillance system and included sociodemographic characteristics (gender, address), date of symptom onset, self-reported household contact with a COVID-19 suspected case (within 14 days, no/yes), personal history of chronic illnesses (no/yes; obesity, asthma, pulmonary obstructive disease, diabetes mellitus, kidney disease, immunosuppression, or human immunodeficiency virus [HIV] infection), disease severity (severe illness defined as dyspnea requiring hospital admission, no/yes), and history of non-communicable diseases. Immunosuppression referred to any identified cause (e.g., malnutrition) of the related deficiency except for that previously cited (personal history of diabetes mellitus, HIV, chronic kidney disease, or asthma). The medical files from the patients are the primary data source of the surveillance system.

The National Urban System database of Mexico,[8] which includes data from 401 cities, was used to classify the place of residence of the enrolled subjects as rural or urban.

Laboratory methods

Clinical specimens (nasopharyngeal or deep nasal swabs) were used, and nucleic acids were extracted from 200 μL of the sample using the MagNa Pure LC Total Nucleic Acid Isolation Kit automated system (catalog: 03038505001; Roche Diagnostics, Mannheim, Germany), as previously described in the study by Fernandes-Matano et al.[9] SARS-CoV-2 detection was performed by using the primers and probes proposed by Corman et al[10] using the Super-Script III Platinum One-step qRT-PCR System (catalog: 12574035; Invitrogen Carlsbad, California, USA) in the 7500 Fast Real-Time PCR System (Applied Biosystems, Foster City, California, EUA).[11]

The analytical procedure was performed in the network of laboratories for epidemiological surveillance of the IMSS.

Statistical analysis

Summary statistics were computed. Prevalence odds ratios (ORs) and 95% confidence intervals (CIs) were used to evaluate factors predicting a positive qRT-PCR result. Bivariate unconditional logistic regression models were used, and a multiple model was fitted.

Ethical considerations

The written informed consent was provided by any parent or legal guardian from the enrolled pediatric patients. This study was approved by the Local Research Ethics Committee (601) of the Mexican Institute of Social Security (approval R-2020-601-015; April 30, 2020).

Results

Data from 1849 individuals were analyzed, and COVID-19 was confirmed in 15.9% ($n = 294$) of them. The study profile is shown in Fig. 1. Most of participants were males (53.3%) and were aged younger than 3 years (43.6%) at time of acute symptom onset (Table 1). Forty-two percent of the analyzed children and teenagers were ambulatory cases, and severe illness was documented in nearly 38% of the subjects.

When compared with participants who tested negative (Table 1), patients with laboratory-confirmed COVID-19 were older (6.5 ± 5.7 vs. 5.3 ± 5.2 years old, $P < 0.001$), were more likely to reside in localities with a population size lower than 15,000 inhabitants (48.0% vs. 38.2%, $P < 0.001$), and showed a higher obesity prevalence (5.8% vs. 2.6%, $P = 0.003$). Discarded cases of COVID-19 were also more likely to require hospital entry (59.7% vs. 48.3%, $P < 0.001$).

Confirmed cases also showed a higher prevalence of self-reported household contact with a case (29.6% vs. 14.1%, $P < 0.002$) within 14 days before acute illness. Sudden disease onset ($P = 0.006$) and milder symptoms ($P < 0.001$) were more frequent among discarded SARS-CoV-2 cases (Table 2). No gender-related differences were observed.

In the multiple regression model (Table 2), a 2-fold increase in the odds of testing positive for COVID-19 was observed among older (13–15 years old) participants (reference: < 3 years old; OR = 2.08, 95% CI = 1.46–2.96), among those with obesity (OR = 2.05, 95% CI = 1.11–3.79), and among children and adolescents with self-reported household contact with a case (OR = 2.27, 95% CI = 1.68–3.08). When compared with localities with low population sizes (lower than 15,000 inhabitants), individuals residing in more crowded locations were less likely to obtain a positive result. Non-severe respiratory symptoms also reduced the odds of laboratory-confirmed COVID-19 in infants and teenagers.

Discussion

Our study characterized factors associated with the odds of laboratory-confirmed SARS-CoV-2 infection in pediatric patients with acute upper or lower respiratory symptoms of infection. The presented results may be useful to identify children and teenagers at increased risk of COVID-19 in whom timely quarantining may reduce viral spread, particularly after the restart of in-person education activities.

Given that the COVID-19 pandemic has been distinguished by a low incidence among children, the strengths of this study include a large number of infants and teenagers enrolled and their national representativeness because the participants were obtained from a nationwide prospective cohort. The use of qRT-PCR, the gold standard of SARS-CoV-2 diagnosis, for clinical specimens from all analyzed subjects is another strength of this study.

As observed in older subjects, we documented an association between increasing age and the odds of a positive test result ($OR_{per\ year} = 1.06$; 95% CI = 1.03–1.08). In the age-stratified analysis, this association seemed to be determined by teenagers because (when compared with infants younger than 3 years) a 2-fold increase in the odds of confirmed SARS-CoV-2 infection was documented among them (OR = 2.04, 95% CI = 1.36–3.05). Differences in the severity of COVID-19 symptoms among older patients result from angiotensin-converting enzyme 2 (ACE2) expression, lymphocyte count, and trained immunity might plan an important role in the observed scenario.[12] Changes in ACE2 activity during puberty have been documented.[13]

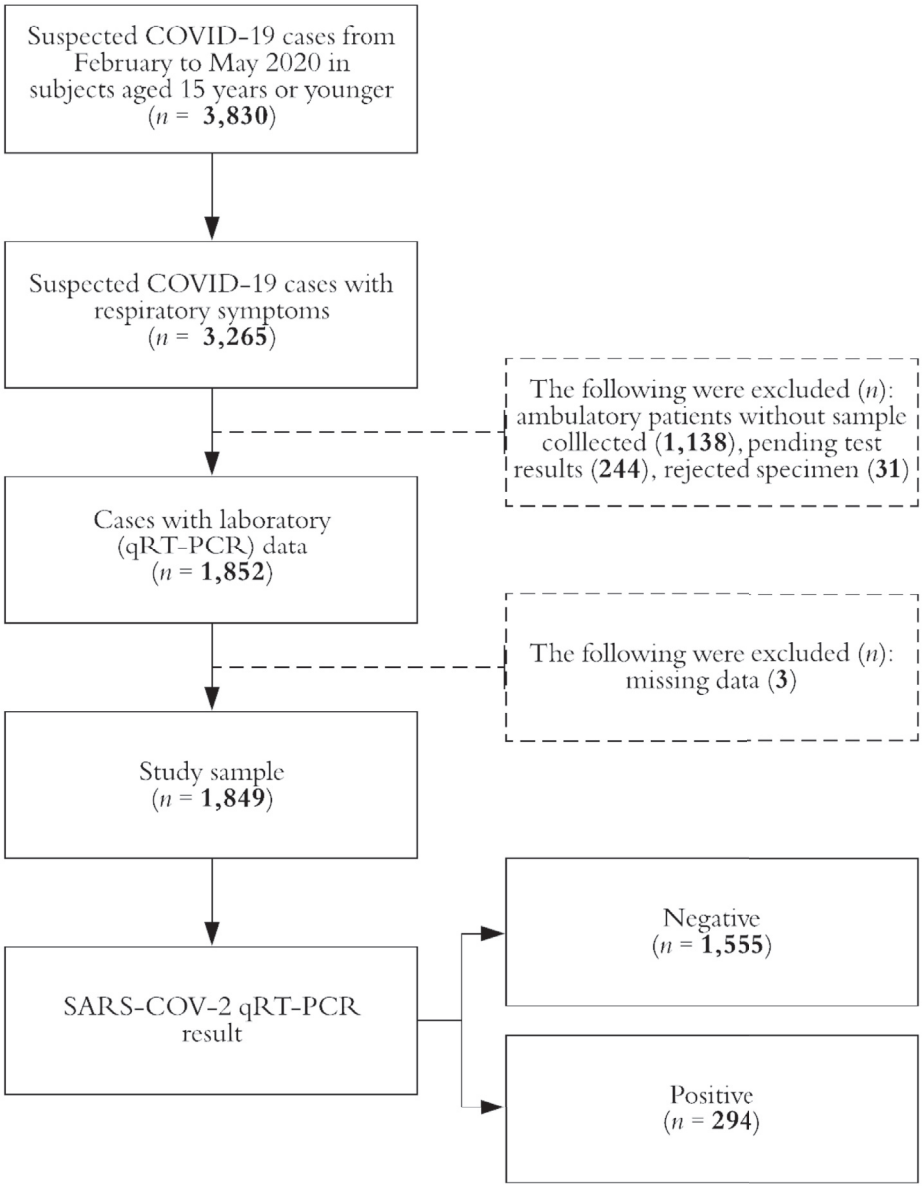

Fig. 1. Study profile, Mexico 2020. COVID-19; coronavirus disease 2019; SARS-CoV-2; severe acute respiratory syndrome coronavirus 2; qRT-PCR, reverse transcription-quantitative polymerase chain reaction.

No gender-related differences in the outcome of interest were observed in the study sample. In male adults, higher disease severity and mortality risk have been documented.[14] The protective role of estrogen among postpubertal women seems to be determining the observed scenario.[15]

After adjustment by age, gender, and other clinical characteristics, participants with obesity were more likely to have a positive qRT-PCR result (OR = 2.46, 95% CI = 1.29–4.71). Obesity has been consistently associated with more severe COVID-19 manifestations,[16] possibly through inflammatory and metabolic pathways.[17] Childhood obesity is also characterized by a low-grade inflammation status,[18] which makes it plausible that children infected by SARS-CoV-2 and with high adiposity levels were more likely to develop respiratory symptoms and to be studied as suspected COVID-19 cases. Globally, Mexico has one of the highest prevalence for children who are overweight or obese, and increasing trends have been documented.[19] The COVID-19 lockdown may worsen the childhood obesity pandemic.[20] However, further research is needed to elucidate the underlying mechanism of obesity-related susceptibility to coronavirus infections.

Nearly 11% of confirmed cases were reported among newborns (mean age = 12.3 ± 8.9 days); however, current data suggest that viral transmission may be secondary to household contact with other cases rather than vertical (intrauterine) or peripartum transmission or through breastfeeding.[21] Moreover, in our research, self-reported household contact with a case was associated with the highest increase in the odds of laboratory-positive COVID-19 (OR = 2.27, 95% CI = 1.68–3.08). Even if controversial, the current

Table 1
Characteristics of the study sample, Mexico 2020.

Characteristic	Overall	SARS-CoV-2 test result		*P*
		Negative	Positive	
	$n = 1,849$	$n = 1,555$	$n = 294$	
Gender				
Girl	863 (46.7)	728 (46.8)	135 (45.9)	0.777
Boy	986 (53.3)	827 (53.2)	159 (54.1)	
Age (years)[a]	5.3 ± 5.2	5.1 ± 5.0	6.5 ± 5.7	<0.001
Age-group (years)				
<3	806 (43.6)	701 (45.1)	105 (35.7)	<0.01
3–5	291 (15.7)	248 (16.0)	43 (14.6)	
6–12	479 (25.9)	404 (26.0)	75 (25.5)	
13–15	273 (14.8)	202 (12.9)	71 (14.2)	
Population by place of residence (×1000)				
<15	735 (39.8)	594 (38.2)	141 (48.0)	<0.001
15–49.9	296 (16.0)	274 (17.6)	22 (7.5)	
50–99.9	94 (5.1)	87 (5.6)	7 (2.4)	
≥100	724 (39.1)	600 (38.6)	124 (42.2)	
Flu vaccinated[b]				
No	1552 (83.9)	1305 (83.9)	247 (84.0)	0.969
Yes	297 (16.1)	250 (16.1)	47 (16.0)	
Household contact with a case[c]				
No	1542 (83.4)	1335 (85.9)	207 (70.4)	<0.001
Yes	307 (16.6)	220 (14.1)	87 (29.6)	
Sudden symptom onset				
No	1211 (65.5)	998 (64.2)	213 (72.5)	0.006
Yes	638 (34.5)	557 (35.8)	81 (27.5)	
Disease severity[d]				
Mild to moderate	1148 (62.1)	932 (59.9)	216 (73.5)	<0.01
Severe	701 (37.9)	623 (40.1)	78 (26.5)	
Personal history of				
Chronic illness (any, yes)	315 (17.0)	264 (17.0)	51 (17.4)	0.877
Obesity (yes)	57 (3.1)	40 (2.6)	17 (5.8)	0.003
Asthma (yes)	104 (5.6)	93 (6.0)	11 (3.7)	0.126
COPD (yes)	2 (0.1)	2 (0.1)	0 (0)	0.538
Diabetes mellitus (yes)	13 (0.7)	10 (0.6)	3 (1.0)	0.478
CKD (yes)	27 (1.5)	21 (1.4)	6 (2.0)	0.366
Immunosuppression (yes)[e]	159 (8.6)	132 (8.5)	27 (9.2)	0.697
HIV (yes)	3 (0.2)	3 (0.2)	0 (0)	0.451

SARS-CoV-2, severe acute respiratory syndrome coronavirus 2; COPD; chronic obstructive pulmonary disease; CKD, chronic kidney disease; HIV, human immunodeficiency virus.
The absolute and relative (%) frequencies are presented, except if other is specified; *P-value* from chi-squared or t-tests as corresponding.
[a] The arithmetic mean ± standard deviation is presented.
[b] During the flu season 2019––20.
[c] Self-reported; within 14 days before the symptoms onset.
[d] Severe illness was defined by dyspnea requiring hospital admission.
[e] Immunosuppression was referred to any identified cause of the related deficiency except for the personal history of diabetes mellitus, HIV, CKD, or asthma.

Table 2
Predictors of a positive SARS-CoV-2 test result, Mexico 2020.

Characteristic	OR (95% CI), *p*					
	Unadjusted			Adjusted		
Gender (Ref: girl)						
Boy	1.04	(0.81–1.33)	0.777	1.08	(0.83–1.40	0.560
Age-group (Ref: <3 years)						
3–5	1.16	(0.79–1.70)	0.458	1.04	(0.70–1.55)	0.841
6–12	1.24	(0.90–1.71)	0.184	1.10	(0.78–1.53)	0.592
13–15	2.35	(1.68–3.30)	<0.001	2.08	(1.46–2.96)	<0.001
Population by place of residence (× 1000; Ref: <15)						
15–49.9	0.34	(0.21–0.54)	<0.001	0.34	(0.21–0.55)	<0.001
50–99.9	0.34	(0.15–0.75)	0.007	0.34	(0.15–0.76)	0.008
100	0.87	(0.67–1.14)	0.309	0.82	(0.62–1.08)	0.154
Household contact with a case (Ref: no)						
Yes	2.55	(1.91–3.40)	<0.001	2.27	(1.68–3.08)	<0.001
Sudden symptom onset (Ref: no)						
Yes	0.68	(0.52–0.90)	0.006	0.80	(0.60–1.06)	0.122
Disease severity (Ref: mild to moderate)						
Severe	0.54	(0.41–0.71)	<0.001	0.69	(0.51–0.92)	0.012
Obesity (Ref: no)						
Yes	2.33	(1.30–4.16)	0.004	2.05	(1.11–3.79)	0.022
Asthma (Ref: no)						
Yes	0.61	(0.32–1.16)	0.130	0.55	(0.29–1.06)	0.075
Immunosuppression (Ref: no)						
Yes	1.09	(0.71–1.68)	0.697	1.22	(0.79–1.91)	0.387

SARS-CoV-2; severe acute respiratory syndrome coronavirus 2; Ref, reference; OR, odds ratio; CI, confidence interval.
Logistic regression models were used to estimate OR and 95% CI. The listed variables were used to obtain the adjusted estimates. Severe illness was defined by dyspnea requiring hospital admission. Immunosuppression was referred to any identified cause of the related deficiency except for the personal history of diabetes mellitus, human virus immunodeficiency infection, chronic kidney disease, or asthma.

consensus regarding the average number of new infections caused by each patient (reproduction number [R]) is about 3.[22]

As per our findings, pediatric patients residing in larger localities were less likely to be diagnosed with COVID-19. The documented association deviates from the World Health Organization COVID-19 preparedness guidance, which states a higher transmission risk in more crowded areas.[23] The research group suggests that this may be secondary, at least partially, to socio-economic gaps between rural and urban areas of Mexico[24] and to lower physical restrictions during the pandemic in less urbanized areas. If later replicated, further research is needed to identify factors determining these findings.

The inclusion of only patients who sought healthcare attention is a limitation of the study and may be implied, among others, in the high documented frequency of severe illness (26.5% and 40.1% in confirmed and discarded COVID-19 cases, respectively). However, and since no mass SARS-CoV-2 screening has been performed in Mexico,[7] we consider that our results are still useful to detect pediatric patients with acute respiratory symptoms who are at increased risk of being positive cases of COVID-19.

Conclusion

Children and teenagers seem to account for a relatively small proportion of COVID-19 cases. However, they may play a role in the spread of respiratory viruses, and their timely identification and isolation may be useful to reduce the related disease burden. We identified factors associated with the odds of laboratory-confirmed disease in a large sample of subjects, and our findings may be useful from a public health perspective, mainly after in-personal scholar activities are reinitialized.

Author statements

Ethical approval

This study was approved by the Local Research Ethics Committee (601) of the Mexican Institute of Social Security (approval R-2020-601-015; April 30, 2020).

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None to declare.

Competing interests

None to declare.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2020.10.012.

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Review Paper

Psychological impact of infectious disease outbreaks on pregnant women: rapid evidence review

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ABSTRACT

Objectives: Infectious disease outbreaks can be distressing for everyone, especially those deemed to be particularly vulnerable such as pregnant women, who have been named a high-risk group in the current COVID-19 pandemic. This paper aimed to summarise existing literature on the psychological impact of infectious disease outbreaks on women who were pregnant at the time of the outbreak.
Study design: The design of this study is a rapid review.
Methods: Five databases were searched for relevant literature, and main findings were extracted.
Results: Thirteen articles were included in the review. The following themes were identified: negative emotional states; living with uncertainty; concerns about infection; concerns about and uptake of prophylaxis or treatment; disrupted routines; non-pharmaceutical protective behaviours; social support; financial and occupational concerns; disrupted expectations of birth, prenatal care and postnatal care and sources of information.
Conclusions: Pregnant women have unique needs during infectious disease outbreaks and could benefit from up-to-date, consistent information and guidance; appropriate support and advice from healthcare professionals, particularly with regards to the risks and benefits of prophylaxis and treatment; virtual support groups and designating locations or staff specifically for pregnant women.

Introduction

The outbreak of coronavirus COVID-19 has – as of June 2nd, 2020 – seen more than six million cases and more than 377,000 related deaths worldwide.[1] Such outbreaks are understandably distressing; there is a wealth of research to suggest a substantial negative psychological impact of public health emergencies such as pandemics.[2] The importance of addressing the negative impact of the COVID-19 pandemic on mental health is reflected in the provision of public guidance for mental health and well-being during the pandemic[3] and calls for psychosocial support to be incorporated into pandemic healthcare; after all, there is 'no health without mental health'.[4]

There is historical evidence of pregnant women being a high-risk group during pandemics; pregnancy was associated with high mortality rates during the H1N1 'swine flu' pandemic[5] and the severe acute respiratory syndrome (SARS) pandemic.[6] During the COVID-19 outbreak, questions have been raised about the particular risks for pregnant women and their unborn babies,[7] with a review of the limited data available so far,[8] suggesting that outcomes for mothers are more promising than for mothers of previous outbreaks. However, on March 16th 2020, the UK government announced that pregnant women were considered a 'vulnerable group' and recommended they self-isolate.[9]

Despite the considerable body of literature on clinical outcomes of being diagnosed with an infectious disease during pregnancy, little attention has been paid to the psychological impact of such outbreaks on pregnant women (including non-infected individuals). They may have fears for their own health, given the physiological changes that occur during pregnancy which may make them more severely affected by infectious diseases,[10] as well as the health of their unborn babies. They may also experience distress due to disrupted prenatal care and delivery: women who gave birth during the SARS and H1N1 outbreaks were discharged as soon as possible after delivery, and prenatal services considered non-essential were suspended.[6,11,12] Research suggests lack of

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control over decisions relating to childbirth can be traumatic,[13] raising concerns about how women giving birth during the COVID-19 pandemic will cope with, for example, restrictions on hospital visitation procedures during and after labour. In many countries (such as the UK), women are being requested to attend all prenatal appointments alone[14] and in some countries (including Poland and China) are required to give birth alone,[15,16] despite familial support during the birthing process being considered essential for women's well-being.[17]

Maternal mental health research is an important, yet understudied, aspect of public health research, as psychological difficulties during pregnancy can affect the mental health of both mothers and children. Buekens et al.[18] have called for research on pregnancy during the COVID-19 pandemic, recognising its potential psychological and social impact. This review is the first to systematically examine literature on the psychological impact of infectious disease outbreaks on pregnant women and factors associated with this impact.

Methods

This was a rapid evidence review in response to the 2019–2020 COVID-19 outbreak. Rapid reviews follow the general guidelines for traditional systematic reviews but are simplified to produce evidence quickly; they are recommended during circumstances where policymakers urgently need evidence synthesis to inform public health guidelines.[19]

Medline, PsycInfo, Embase, Global Health and Web of Science databases were searched from inception to the date of the searches (April 1st 2020). The search involved a combination of pregnancy-related terms (e.g. pregnant, pregnanc*), mental health–related terms (e.g. mental health, well-being) and outbreak-related terms (e.g. pandemic, SARS). The full search strategy is presented in Appendix I.

Inclusion criteria for the review were as follows: articles must (i) report primary data either quantitative or qualitative, (ii) be published in peer-reviewed journals, (iii) be written in English and (iv) report on psychological effects of emerging infectious disease outbreaks (e.g. SARS, H1N1) on women who were pregnant at the time of the outbreak. Zika was also considered a relevant outbreak for review because, despite the spread of the infection being different, this outbreak had a particular impact on pregnant women as it was associated with birth defects.[20] The review focused on emerging infectious diseases (that is, diseases appearing for the first time in a population or rapidly increasing in incidence or geographic range).[21] Articles were excluded if they were letters, commentaries or reviews without primary data; if they were written in any language other than English; if they did not consider a psychological aspect of being pregnant during an infectious disease outbreak and if the disease outbreak was not an emerging infectious disease (i.e. articles on seasonal influenza were excluded).

The authors ran the searches and downloaded all citations to EndNote version X9 (Thomson Reuters, New York, United States) where duplicates were removed. Titles and then abstracts were screened by one author (SKB) for relevance to the selection criteria. Full texts of all articles remaining after abstract screening were downloaded and assessed to decide whether they met all inclusion criteria. Reference lists of all included articles were hand searched, and any references not found by our own search, which suggested from their title that they may contain relevant data, were downloaded and assessed for eligibility. Screening was performed by one author; however, any uncertainties about whether an article met all inclusion criteria or not were discussed with the other authors.

Spreadsheets were created to systematically extract the following data from the articles: country, design, infectious disease outbreak, participant information (n and sociodemographics), measures used and results. During data extraction, it became clear that the included studies used varying methodologies and measures and presented their data in various ways. Therefore, study results were synthesised using inductive thematic analysis to code the data and organise into themes.[22] This was chosen as an effective way of describing data from multiple studies. Following repeated readings of the 'results' section of our data spreadsheet, data were first broadly coded and used to develop descriptive themes. Similar results were grouped together (e.g. all data on anxiety, stress or fear

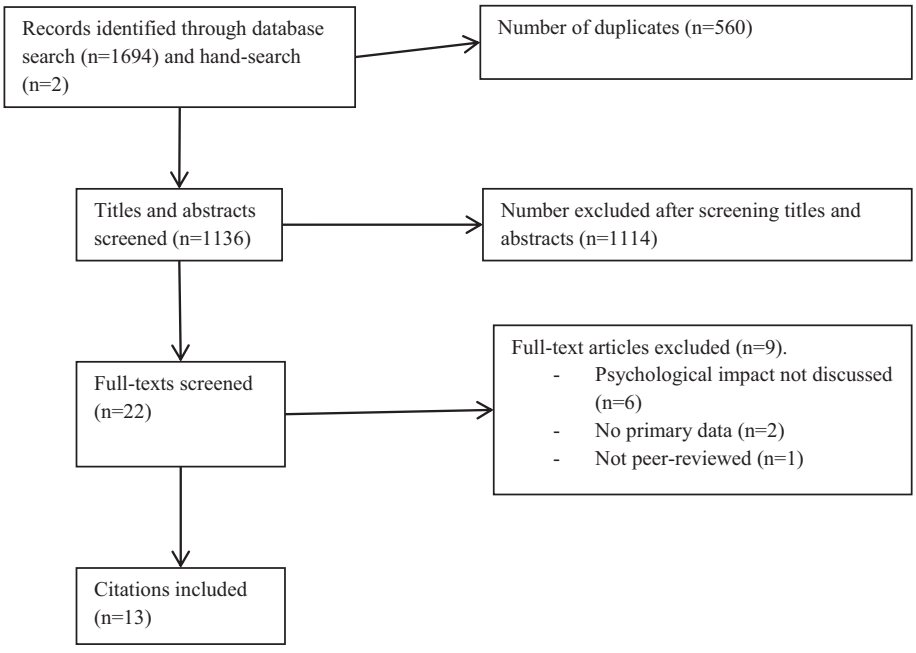

Fig. 1. . Flow diagram of the screening process.

Table 1
Study characteristics of included articles.

Study	Country	Disease outbreak	Participants	Measures
Dodgson et al. (2010)[23]	China (Hong Kong)	SARS	8 women who delivered healthy babies during the outbreak; mean age 34.3 years (range 28–38)	Interviews about experiences of being pregnant and delivering their baby during the SARS epidemic
Lee et al. (2006)[24]	China (Hong Kong)	SARS	235 women pregnant during the outbreak compared with a historical cohort of 939 recruited a year before; mean age 29.9 years (SARS cohort), 29.6 years (pre-SARS cohort)	Beck Depression Inventory, Spielberger State-Trait Anxiety Inventory, Medical Outcomes Study Social Support Survey. The SARS cohort also completed a 41-item questionnaire on worries, perceived risk and behavioural responses to SARS
Linde & Siqueira (2018)[26]	Brazil, Puerto Rico and USA	Zika	18 women: 5 had a recently born baby, 6 were pregnant, 5 were planning to get pregnant, 3 had no plans to get pregnant. Age range 22–41	Interviews about personal and family life, perceptions and knowledge of Zika, views on reproductive health and rights regarding the Zika syndrome
Lohm et al. (2014)[27]	Australia and Scotland	H1N1	14 pregnant women aged between 20 and 40 years	Interviews and focus groups about experiences with H1N1 and the public health response to H1N1
Lyerly et al. (2012)[28]	USA	H1N1	22 pregnant women who had participated in the H1N1 vaccine trials; mean age 31 years, range 19–39	Interviews about experiences of decision-making around participation in the H1N1 vaccine trial
Lynch et al. (2012)[29]	USA	H1N1	144 women: 43.4% of women were pregnant and 56.6% were within 6 months postpartum; 26.4% aged 18–24 years, 61.8% aged 25–34 years and 11.8% aged 35–44 years	Focus groups covering perceptions and awareness of H1N1, influenza vaccinations and antiviral medicines and trusted sources of information
Meireles et al. (2017)[31]	Brazil	Zika	14 pregnant women: 6 in the first trimester, 5 in the second trimester and 3 in the third trimester; mean age 33.4 years, range 28–40	Focus groups with questions on feelings and experiences around being pregnant during the Zika outbreak
Ng et al. (2013)[25]	China (Hong Kong)	SARS	980 pregnant women of at least 16 weeks gestation; 0.6% aged younger than 18 years, 80.7% aged 18–35 years and 18.7% aged older than 35 years	Study-specific survey asking about sociodemographics, SARS knowledge, socio-economic impact of SARS and Chinese version of the State-Trait Anxiety Inventory
Ozer et al. (2010)[32]	Turkey	H1N1	314 pregnant women; 27.4% in the first trimester, 33.8% in the second trimester and 38.8% in the third trimester	48-question study-specific survey covering vaccination status, factors affecting decisions about vaccinating, H1N1 vaccine side-effects and beliefs about H1N1 vaccination campaign conspiracy
Sakaguchi et al. (2011)[33]	Canada	H1N1	130 pregnant women who called counselling service Motherisk for counselling regarding the safety of H1N1 vaccine; median age 33 years, range 21–45; 31.5% in the first trimester at time of call, 46.2% in the second trimester, 22.3% in the third trimester	Study-specific questionnaire including questions on vaccination status, decision-making and factors that precipitated call to Motherisk
Sasaki et al. (2013)[34]	Japan	H1N1	109 pregnant women attending prenatal classes	Study-specific questionnaire measuring anxiety, satisfaction with information supplied, reasons for anxiety, prophylaxis interventions practiced
Sim et al. (2011)[35]	Scotland and Poland	H1N1	10 pregnant women	Interviews covering socio-economic background, migration history, family circumstances, general health during pregnancy, views of healthcare received during pregnancy, perceptions and experience of H1N1 influenza and the vaccine, sources of information about H1N1 and the vaccine, government responses to the pandemic and decision-making about the H1N1 vaccine
Steelfisher et al. (2011)[30]	USA	H1N1	514 pregnant women	Study-specific survey with approximately 84 questions relating to attitudes and experiences associated with the H1N1 vaccine

SARS, severe acute respiratory syndrome.

were grouped together and coded as 'negative emotional states'; all data on work or money concerns were grouped together and coded as 'financial and occupational concerns'). The final list of themes was reached when no new themes emerged from the data.

Results

Initial searches yielded 1694 citations; 560 duplicates were removed. Nine hundred ninety-nine articles were excluded based on title, 115 excluded based on abstract and nine excluded after assessing full texts. Two additional articles were found via hand searching, leaving a total of thirteen articles included (refer Fig. 1 for flow diagram).

Study characteristics

Studies were international, including participants from China;[23–25] Brazil, Puerto Rico and the USA;[26] Scotland and Australia;[27] the USA;[28–30] Brazil;[31] Turkey;[32] Canada;[33] Japan[34] and Scotland and Poland.[35] Outbreaks included SARS (n = 3), H1N1 (n = 8) and Zika (n = 2). Articles were published between 2006 and 2018. The number of participants ranged from 8 to 980. A variety of

quantitative and qualitative measures were used by the studies, with more than half (n = 7) using interviews or focus groups. Study characteristics are presented in more depth in Table 1.

A detailed summary of the thematic analysis is presented in Table 2. It should be noted there is some overlap between themes; specifically, 'negative emotional states' in general emerged as a theme, but many articles also referred to negative emotions due to specific factors, which have also been classified as themes (for example, 'concern about risk of infection' is a negative emotional state but also a theme in itself).

Negative emotional states

State anxiety was significantly higher in pregnant women during the SARS pandemic than in a comparative pre-SARS group in China,[24] while another Chinese study during the SARS pandemic found that 22.6% of 980 participants reported high anxiety and 65.2% moderate anxiety.[25] Other negative emotions reported across different outbreaks and in different countries included sadness, uneasiness, fear, panic, tension, loss of control of life, shame, failure and guilt due to the pressure of having a healthy child;[26] unease even when at home, feeling a lack of security and a loss of freedom;[24] stress;[27] frustration, anxiety and sleep problems;[23] pandemic-related anxiety;[34] pressure from others regarding infection prevention[31] and negative body image due to wearing protective clothing.[31]

Living with uncertainty

Participants in various countries during the SARS, Zika and H1N1 outbreaks reported living with uncertainty, mostly due to doubt and confusion about the risk to their health and that of their baby.[23,26,27,31] Uncertainty was worsened by conflicting and rapidly changing media messages and not receiving recommendations from doctors regarding what mothers should and should not be doing during pregnancy, according to a SARS-related interview study from Hong Kong.[23]

Concerns about infection

Participants in various countries and experiencing different outbreaks expressed concerns about the health of themselves and their babies,[24,25,27,29,31,30] including fears that infection could lead to miscarriage or preterm delivery.[24] Concerns about infection appeared to depend on perceptions of risk: many participants in a large Chinese study of the SARS pandemic overestimated their risk of being infected[25] while others in the USA during the H1N1 outbreak reported during focus groups that they believed they were less vulnerable as they believed pregnant women had stronger immune systems due to prenatal vitamins, healthy diet and exercise.[29]

Concerns about, and uptake of, prophylaxis/treatment

Pregnant women across many of the countries and outbreaks studied expressed concerns about antivirals[24,29] and vaccinations,[27,29,30,32,33,35] mostly due to potential side-effects for the developing foetus.[29,30,32,33,35] Reasons for lack of uptake of vaccines included anticipating changing knowledge of side-effects,[27] thinking it unnecessary[33] and previous adverse vaccination effects.[33] Many participants in the USA during the H1N1 outbreak reported being cautious about taking any medications at all during pregnancy;[29] others in Scotland and Poland during the same outbreak noted during interviews the contradiction between the culture of caution characterising most pregnancy-related health advice and being urged to have a relatively untested vaccine.[35] Conversely, motivators for receiving vaccines included concerns about infants' well-being,[28,29,33,35] previous history of complication or illness from influenza,[33] having contact with vulnerable people[33] and knowledge about the minimal risks of vaccination.[32]

Disrupted routines

Pregnant women's daily routines, social lives and leisure activities were disrupted as they tried to eliminate the risk of contracting the diseases.[23–26] Some participants in China did not leave their homes at all during the SARS pandemic[23,24] which led to feeling confined especially when living in a small apartment.[23] Relationships with spouses were affected due to decreased intimate contact[25] and sleeping separately due to fear of infection.[23]

Non-pharmaceutical protective behaviours

Participants across different countries and outbreaks reported living in a state of vigilance related to hygiene measures and adopting new behaviours to mitigate their risk of contracting infection such as monitoring the news and information gathering;[23,34] avoiding places of risk;[23,29,31] gathering hygiene supplies;[23] cleaning hands vigilantly;[23,24,29,34] washing bags, clothes and hair after leaving the house;[23] wearing masks;[24,25,29,34] stocking up on prophylaxis materials[34] and cancelling planned visits from family or banning visitors to the home altogether.[23] During the Zika outbreak, participants reported using insect repellents constantly and wearing long sleeves and closed shoes which caused discomfort.[26]

Social support

Women pregnant during the SARS outbreak reported significantly higher affectionate support, positive social interaction and informational support than a pre-SARS cohort.[24] Social support appeared to mediate symptoms of depression; the authors noted a significant negative correlation between depression and social support.

Financial and occupational concerns

In one study of SARS in Hong Kong,[25] more than a third of participants reported their family's financial situation had been negatively affected by the outbreak. Participants reported increased expenses due to using taxis because buses and subways were considered unsafe[23] and having to buy supplies to mitigate their risk of infection, such as masks and cleaning supplies, during the SARS pandemic[23] or insect repellents and clothing during the Zika outbreak.[31] Some participants took early maternity leave and forfeited pay if they worked in high-risk occupations such as healthcare[23] or made special leave arrangements;[25] others risked their careers by giving up career-promoting opportunities which involved attending meetings or travelling.[26]

Table 2
Themes emerging from included studies.

Theme	Reference	Evidence
Negative emotional states	Dodgson et al. (2010)[23]	Participants reported frustration, anxiety and difficulty sleeping.
	Lee et al. (2006)[24]	State anxiety was higher in pregnant women (mean score 37.2) during the SARS pandemic than in a comparative pre-SARS group (mean score 35.5, $P = 0.02$) while no significant difference was found in trait anxiety scores. The SARS cohort was slightly more likely to score highly on depression but not significantly. Among all, 18.4% of women felt uneasy even at home due to SARS, 54.7% felt a lack of security and 48.3% felt a loss of freedom.
		Participants reported worries and fears, primarily regarding the risk of infection (refer 'concerns about risk of infection' theme).
	Linde & Siqueira (2018)[26]	Participants reported sadness, uneasiness, fear, helplessness, panic, tension, responsibility, shame, failure, guilt due to pressure of having a healthy child, perceived loss of control of their own lives.
	Lohm et al. (2014)[27]	Participants reported emotional stress.
	Meireles et al. (2017)[31]	Participants reported a negative impact on body image due to not being able to show their bump or wear dresses that emphasised their pregnancy and having to cover up in clothing that made them feel constrained. Participants felt that others (e.g. their partners and parents) placed demands of them regarding prevention of Zika, leaving them feeling under pressure.
		Participants reported anxiety around the impact of the virus – refer 'concerns about risk of infection' subtheme.
	Ng et al. (2013)[25]	The mean state anxiety score (measured by the State-Trait Anxiety Inventory) was 50.4 (range 23–80). Among all, 65.2% of participants experienced moderate anxiety, 22.6% high anxiety and 12.2% low-level anxiety. Age, marital status, gestational age, parity, education level and gestational complications were not significantly associated with anxiety level, but there was a significant relationship between the state anxiety score and extent of socio-economic impact ($P < 0.01$) where higher anxiety was associated with higher socio-economic impact.
	Sasaki et al. (2013)[34]	Among all, 96.3% of participants felt concerned or strongly concerned about the pandemic. Nearly, all who felt anxious cited their pregnancy as the main reason for this.
Living with uncertainty	Dodgson et al. (2010)[23]	Participants reported doubt and confusion about what was a true threat to themselves and their babies, often due to conflicting and constantly changing messages in the media. All reported receiving no recommendations from doctors regarding what they should and should not be doing during pregnancy and postpartum; all but two found this frustrating and said it added to their anxiety about their baby's safety.
	Linde & Siqueira (2018)[26]	Participants reported uncertainty and mistrust concerning unknown factors surrounding Zika, contributing to feelings of helplessness and distress.
	Lohm et al. (2014)[27]	Participants reported that the unknown effects of both infection and vaccination against infection increased their emotional stress.
	Meireles et al. (2017)[31]	Participants reported uncertainty about the impact of the virus.
Concerns about infection	Lee et al. (2006)[24]	Pregnant women tended to overestimate their risk of contracting SARS: 21.9% of participants believed they were likely or very likely to contract it, while 21.5% believed their newborns were likely to.
		Almost half (49.6%) of their participants were worried or very worried about contracting SARS themselves, while 58.1% of participants were worried or very worried about their newborn contracting it; 63.2% of participants were worried about their spouse contracting it; and 57.3% of participants were worried about relatives or friends contracting it.
		Among all, 46.6% were worried or very worried about infection leading to miscarriage, and 46.2% of participants were worried or very worried about infection leading to preterm delivery.
		Fears could lead to disrupted healthcare: 66.7% of participants feared antenatal visits in the hospital, 79.9% feared any consultations in the hospital, 12.0% cancelled appointments in the hospital and 38.9% considered doing so and 20.9% postponed appointments in the hospital while 29.1% considered doing so.
	Lohm et al. (2014)[27]	Participants reported concerns about the health of both themselves and their babies.
		Not knowing anyone who had the virus provided an impression of safety, and many were not too worried if they did not know of any cases. By contrast, others were mobilised into action (such as stopping work, pulling children out of school and no longer leaving the house) if the pandemic broke out in their neighbourhood.
	Lynch et al. (2012)[29]	Participants did not show high levels of concern: 25.2% of participants were not at all worried, and many doubted the outbreak was as severe as reported and blamed the media for generating mass hysteria. Although many did not initially perceive H1N1 to be severe or personally threatening, views shifted during group discussions and exposure to news media and raised levels of concern.
		Concerns about infection appeared to depend on perceptions of risk: some participants reported awareness that pregnant women were at a higher risk for H1N1 and cited pregnancy as the main reason for their concern about infection, while others believed they were less vulnerable and pregnant women had stronger immune systems due to prenatal vitamins, healthy diet and exercise. Most reported a limited understanding of the potential severity of H1N1 during pregnancy, and many were confused about how H1N1 differed from seasonal influenza. The primary source of this confusion was lack of consistent messages, particularly from the media.
		Concern about infection was higher among women in cities where H1N1 was most active and lower in cities where the outbreak had not yet peaked.
	Meireles et al. (2017)[31]	Participants reported uncertainty, anxiety and fear around the impact of the virus on both themselves and their baby.
	Ng et al. (2013)[25]	Among all, 71.4% of participants perceived that pregnant women would have a higher risk of being infected. Eighty-nine percent of participants believed their unborn babies would be affected if they contracted it. Ninety-eight percent of participants were worried about getting infected.
	Steelfisher et al. (2011)[30]	Thrity-four percent of participants were concerned they might get sick from H1N1, and 49% of participants were concerned their baby might get sick. Fifty-two percent of participants believed pregnant women were more likely to become seriously ill than the general population from H1N1.
Concerns about, and uptake of, prophylaxis/treatment	Lee et al. (2006)[24]	Among all, 68.8% of participants were worried or very worried about foetal malformation if antiviral drugs were needed for infection.
	Lohm et al. (2014)[27]	Participants reported difficulties in deciding whether to get vaccinated or not; some delayed vaccination due to anticipating changing knowledge of the side-effects.

Table 2 (*continued*)

Theme	Reference	Evidence
	Lyerly et al. (2012)[28]	Participants universally articulated positive or neutral valuation of risks and benefits associated with the H1N1 vaccine (although it must be noted, all participants had taken part in the vaccine trial and therefore are likely to have positive views of the vaccine and are not necessarily representative of pregnant women as a whole). Many believed the risk of contracting H1N1 outweighed any theoretical risk from vaccine. Many jumped at the chance to participate in the trial due to early access to the vaccine. Notions of a growing pandemic and finite supply of vaccine made them eager to have it early, particularly women nearing the end of their pregnancies. Many felt reassured by the research question itself which was focused on dosing rather than vaccine-related harm and made the vaccine seem already safe.
	Lynch et al. (2012)[29]	Women had concerns about both vaccinations and antiviral medicine and were not well informed about either: 41.1% of participants had low acceptance of the H1N1 vaccine, mostly due to concerns about the vaccine being untested and uncertainty about side-effects, particularly long-term side-effects for the developing foetus. Most were unaware of how antivirals work, confusing them with both antibiotics and vaccines, and some were hesitant about potential side-effects of antivirals on their unborn baby. In fact, many were cautious about taking any medications during pregnancy for the same reason. Concern about infant's well-being, however, was a strong motivator for adopting preventive recommendations including vaccination. Among all, 43.5% of participants would take antivirals such as Tamiflu.
	Ozer et al. (2010)[32]	Of all, 8.9% of participants got the H1N1 vaccine. The percentage of participants who felt comfortable with decisions about the vaccine, who did not feel comfortable and who felt hesitant was 68.5%, 7.3% and 24.2%, respectively. Probability of receiving a vaccine was 3.46 times higher among working women than among housewives, 1.85 times higher among women who already had a child and 1.29 times higher among women with a high school education or higher. Correct knowledge about minimal risks associated with vaccine was associated with increase in receiving vaccine. Age, education, place of residence, chronic disease situation and trimester were not significantly associated with vaccination status. Among all, 70.1% of participants believed the vaccine could cause miscarriage, 74.2% thought it could cause deformation in children and 72.3% were worried vaccine could cause infertility.
	Sakaguchi et al. (2011)[33]	Among the 104 participants who received the H1N1 vaccine, concern about risk of H1N1 in foetus and/or themselves was the most cited reason for decision (73.1%), followed by recommendations encouraging vaccination (34.6%) and previous history of complication or illness from influenza (3.8%). More than 20% of participants cited having household contacts (infant or elderly relative) or being a caregiver as contributing to decision. Among those who did not get the vaccine (n = 26), concern about safety of vaccine for themselves and/or foetus was the most cited reason (42.3%) followed by not thinking it necessary (23.1%) and previous adverse events associated with vaccinations (7.7%).
	Sim et al. (2011)[35]	Almost all (9/10) had a critical stance towards H1N1 vaccine. Deciding whether to have the vaccine or not was difficult and anxiety provoking for all and was seen as choosing the 'least worst'' option in terms of competing risks. Participants identified a contradiction between the culture of caution which characterises pregnancy-related advice and being urged to accept a relatively untested vaccine. The risk of being seen as a 'bad mother' for whichever course of action they took heightened the anxiety surrounding decision-making. The unborn baby was the primary concern in weighing up risks and benefits of having the vaccine; the protective effect of the vaccine on the baby was a key motivator, both to protect the baby in utero and also after birth. Participants were concerned about the vaccine being relatively untested, and what was perceived to be a lack of evidence about long-term efficacy and side-effects for both women and unborn babies.
	Steelfisher et al. (2011)[30]	Those who were concerned about their babies getting sick were more likely to have the H1N1 vaccine (50% v 33%), as well as those who believed they themselves were at greater risk than the general population of becoming seriously ill (54% v 28%). Main reasons for not having vaccine: concerns about safety risk to unborn babies (62%) and to themselves (59%); not believing they were at risk of getting H1N1 (15%) or that they would get seriously ill from it (15%); ability to get medication if they did become sick (11%). Sixty-seven percent of participants felt the H1N1 vaccine was safe, compared with 81% who felt the seasonal influenza vaccine was safe for pregnant women. Women who believed it was safe were more likely to get the vaccine (86% v 27%). Sixty-two percent of participants discussed the vaccine with their healthcare provider. Pregnant women who received a recommendation from their healthcare provider to get the vaccine were more likely to have it (65% v 18%).
Disrupted routines	Dodgson et al. (2010)[23]	Daily routines were disrupted, often leading to relationship difficulties with spouses. Examples included sleeping separately from partners if their partner had a high-risk occupation, avoiding contact with other family members, not leaving the house. Not leaving the house left participants who lived in small apartments feeling confined. Participants also did less shopping for food and baby supplies.
	Lee et al. (2006)[24]	Many participants stopped leaving the house.
	Linde & Siqueira (2018)[26]	Participants reported eliminating leisure activities.
	Ng et al. (2013)[25]	Decreased social activities: 4.5% not at all, 32.1% somewhat, 38% moderately, 25.4% very much. Decreased intimate contact with partner: 30.5% not at all, 40.2% somewhat, 22.3% moderately, 7% very much. Decreased social contact with friends: 16.9% not at all, 37% somewhat, 33.9% moderately, 12.3% very much.
Non-pharmaceutical protective behaviours	Dodgson et al. (2010)[23]	All participants reported living in a state of intense vigilance related to hygiene measures. Behaviours included monitoring the news, gathering hygiene supplies, ensuring anyone who entered their homes abided by the current recommendations, cleaning hands vigilantly, washing bags, clothes and hair after going out, cancelling planned visits from family or banning visitors from the home entirely.
	Lee et al. (2006)[24]	Participants reported adopting behavioural strategies to mitigate their risk of contracting infection, including washing hands more than usual (91.5%), wearing masks most or all of the time (70.1%), wearing gloves most or all of the time (1.7%), rarely or never leaving the house (37.2%) and going out less than usual (54.7%).
	Lynch et al. (2012)[29]	Likelihood of taking the following recommendations: 100% of participants would wash their hands and cover coughs; 74.6% would keep children at home; 68.1% would stay away from large gatherings; 43.9% would get alternative prenatal care such as appointments being held over the telephone or at a different location; 36.8% would wear a mask.

(*continued on next page*)

Table 2 (*continued*)

Theme	Reference	Evidence
	Linde & Siqueira (2018)[26]	Participants reported using repellents constantly and wearing long sleeves and closed shoes which often caused discomfort.
	Meireles et al. (2017)[31]	Participants avoided places of risk.
	Ng et al. (2013)[25]	Wearing a mask: 61.2% very much, 25.4% moderately, 10.6% somewhat, 2.8% not at all. Increased personal hygiene: 54.5% very much, 31.1% moderately, 10.8% somewhat, 3.6% not at all. Increased environment disinfection: 46.2% very much, 36.4% moderately, 14.6% somewhat, 2.9% not at all. Increased awareness of infection prevention: 50.8% very much, 37.8% moderately, 9.7% somewhat, 1.6% not at all.
	Sasaki et al. (2013)[34]	Major precautions taken included wearing a mask, stocking up on 'prophylaxis materials' (not clear from article what these were) and information gathering. Nearly all practiced hand washing; other measures included gargling and wearing a mask.
Social support	Lee et al. (2006)[24]	Women who were pregnant during the SARS outbreak reported significantly higher affectionate support ($P = 0.03$), positive social interaction ($P = 0.01$) and informational support ($P = 0.03$) than the pre-SARS cohort, although the groups did not differ on tangible support. Only 10.8% of the SARS cohort reported feeling lonely during the outbreak. Social support appeared to mediate symptoms of depression; the authors noted a significant negative correlation between depression scores and social support scores ($P < 0.0001$).
Financial and occupational concerns	Dodgson et al. (2010)[23]	Some participants took early maternity leave from work with no pay if they worked in high-risk occupations such as healthcare. Other decreases in income were noted due to added expenses of having to use taxis as buses and subways were considered unsafe and having to spend money on masks and cleaning supplies.
	Linde & Siqueira (2018)[26]	Several participants placed careers at risk by giving up growth opportunities such as attending meetings and travelling for work; many tried to work from home or change occupation, often leading them to feel isolated from their colleagues.
	Meireles et al. (2017)[31]	Participants reported additional expenses due to needing to buy repellents and appropriate clothing.
	Ng et al. (2013)[25]	Among all, 24.5% of participants reported somewhat negative socio-economic impact of SARS on daily life, 27.5% moderately, 30.2% very much so, 17.8% not at all. One third stated their family's financial situation had changed. There was a significant relationship between the state anxiety score and extent of socio-economic impact ($P < 0.01$) Some participants made special leave arrangements from work: 43.6% not at all, 24.5% somewhat, 15.5% moderately, 16.4% very much.
Disrupted expectations of birth and prenatal/postnatal care	Dodgson et al. (2010)[23]	None of the women had the birth experience they had hoped for, due to changes in hospital practices. Fifty percent of participants reported that they could not have family members visit them in the hospital; 25% of participants reported that the father was to be the only visitor; 37.5% of participants had restricted time with their own babies as they were kept separately in the hospital nursery. They had to wear masks and gowns and could not kiss their babies, while fathers could only see them through glass, leading to concerns about lack of time for bonding and attachment. There were scheduled feeding times and if they missed one they had to wait for the next. Three participants who had planned deliveries in public hospitals opted instead to pay for private hospitals; participants reported monitoring the visiting policies of their chosen hospitals as well as whether there were SARS cases in those hospitals. One chose a caesarean delivery in a private hospital as her husband would not have been allowed to accompany a natural delivery. One participant reported having to wear a mask during labour which made her sick and caused difficulty breathing. Others reported a lack of pain relief during labour (for example, not being allowed to breathe nitrous oxide to prevent the spreading of disease). Participants reported feeling a lack of connection with healthcare providers in antenatal classes (due to having to sit at the back of the room and nurses all having masks on), as well as minimal contact with medical staff and less than optimal care during delivery. Participants also reported a lack of discharge teaching, so they were sent home not knowing how to properly change nappies or bathe their babies.
Sources of information	Lyerly et al. (2012)[28]	Participants felt they got more detailed information about the H1N1 vaccine from researchers in the vaccine trial than their doctors.
	Lynch et al. (2012)[29]	Highly trusted sources of information were healthcare providers such as obstetricians, midwives and paediatricians and government health agencies; many distrusted the media which they perceived to be benefiting financially from the outbreak, and in some cases, this distrust extended to government officials. Participants preferred the internet or social networks for communication because of immediate access and low cost. Participants with older children also recommended schools as a helpful medium for disseminating information. Most agreed that information should be disseminated in multiple ways through many channels.
	Sakaguchi et al. (2011)[33]	More than 60% of participants reported information from direct healthcare providers or Motherisk was helpful. More than 65% of participants found information from media was confusing and unhelpful.
	Sasaki et al. (2013)[34]	Users of municipality information reported using many more information sources than non-users. Major information sources used were television, internet and newspapers. Nearly all used television; fewer than 30% obtained information from a hospital or clinic, despite being seen regularly for appointments. Many felt that too little information was available.
	Sim et al. (2011)[35]	Participants did not feel official information about H1N1 vaccine addressed concerns in sufficient detail and sought information from a variety of sources. Four women perceived official information about H1N1 vaccine to be a form of propaganda. All sought out alternative information primarily through social networks and the internet. Lack of information about side-effects on unborn baby was the most significant gap in official information.

SARS, severe acute respiratory syndrome.

Disrupted expectations of birth and prenatal/postnatal care

One interview study[23] reported on disrupted expectations of birth, prenatal care and postnatal care in Hong Kong during the SARS outbreak. No participants had the birth experience they had hoped for, due to changes in hospital practices. Prenatal care was also affected: participants reported feeling a lack of connection with healthcare providers in antenatal classes due to having to keep their distance from nurses and nurses having masks on. In terms of postnatal care, participants reported a lack of discharge teaching, so they were sent home not knowing how to properly care for their babies.

Sources of information

Healthcare providers and government health agencies were generally highly trusted as sources of information.[29,33] Many expressed distrust of the media or found it confusing and unhelpful.[29,33] Conversely, in a Japanese quantitative study of the H1N1 outbreak, Sasaki et al.[34] found that television, internet and newspapers were the most common sources of information about the H1N1 outbreak. In the study by Sim et al.,[35] participants in Scotland and Poland did not feel that official information about the H1N1 vaccine addressed their concerns, particularly about potential effects on unborn babies, and sought information from a variety of sources such as social networks and the internet. It is also noteworthy that participants who had taken part in an H1N1 vaccine trial felt they had received more information from the trial's researchers than they had from their doctors.[28]

Discussion

This review suggests disease outbreaks can have a negative emotional impact on pregnant women, creating anxiety, distress and fear which are exacerbated by uncertainty; concerns about infection; concerns about prophylaxis or treatment; disrupted routines; financial and occupational concerns and disrupted expectations of healthcare. Intense vigilance with regards to non-pharmaceutical protective behaviours was frequently reported. Social support may be a protective factor for poor mental health although during an outbreak may be difficult to access. Given the critical role of mental health provision in combatting outbreaks such as COVID-19 and the reciprocal relationship between mental health and physical health,[4] it is important to understand the implications of these findings to help inform public health interventions or campaigns.

While it is likely that outbreaks can cause anxiety for all, one study[24] found that participants of a pre-SARS cohort were less anxious than a group of participants who were pregnant during the SARS outbreak; another study[34] found that, nearly all participants cited that being pregnant during an outbreak was their primary reason for feeling anxious. This is concerning as previous research suggests that experiencing prenatal stress can lead to adverse birth outcomes.[36] Early identification of mental health issues in perinatal patients is essential; midwives should be aware of pregnant women's propensity to experience anxiety during outbreaks and take account of the potential impact of such symptoms on their physical and mental health. Early identification of problems can allow obstetric providers to partner with mental health specialists to establish appropriate treatment plans[37] and provision of public health education and mental health services specifically for pregnant women.[25]

Stress has been frequently linked to uncertainty across the population as a whole[38] but is particularly concerning for pregnant women as previous research suggests that uncertainty can cause fear and distress in pregnancy[39] which could lead to adverse birth outcomes.[36] Public health officials can reduce uncertainty by ensuring that information provided to the public is timely, accurate and consistent with information from other sources. Information directly from healthcare providers and official public health organisations appears preferable. Distrust of media reporting may be prevalent across the population as a whole.[40] The current outbreak advice is not to watch much media and seek information only from trusted sources;[41] pregnant women can take action to avoid media if it causes anxiety.

Many participants expressed concerns about becoming infected, with some overestimating the risk of infection during pregnancy. This highlights the need for timely dissemination of accurate public health information and for clinicians to monitor for overestimation of risk among pregnant women and clear up misconceptions. Where simple advice and reassurance does not work, there may be benefit in brief psychotherapy using a cognitive-behavioural model to reduce anxiety and the associated risk of pregnancy complications.[42,43]

Concerns about prophylaxis or treatment were prevalent, perhaps unsurprisingly as pregnant women have historically low vaccination rates for seasonal influenza[44] and pandemic influenza.[45,46] The decision about whether to receive vaccines or medications may be distressing as pregnancy is already a time when women are faced with cultural expectations of motherhood and any examples of not abiding by advice can lead to women being seen as undisciplined.[47] It is essential that pregnant women are aware of trustworthy, up-to-date information about the risks and benefits of vaccines and medications, particularly given the potential for pregnant women to be identified as a priority group for any vaccination programme as was the case in the UK during the H1N1 pandemic.[48]

Participants reported disrupted routines and changes to relationships with others due to social distancing. This is concerning as social support is essential in enhancing resilience during times of crisis[49] while poor social support is associated with negative psychological outcomes,[50] as is the isolation felt by people quarantined during pandemics.[51] Mental health campaigns aimed at encouraging communication via phone or internet during physical isolation may be useful.[51] Support from others with similar experiences can be particularly helpful to alleviate stress in pregnancy,[52] and social media is a substantial source of support for pregnant women and new mothers[53]; therefore, virtual support groups specifically for pregnant women to support each other may be beneficial. Some of these recommendations (particularly concerning signposting to resources and the use of social media to connect with others) are reflected in existing public health guidance for maintaining good mental health during the COVID-19 pandemic.[3]

Financial and occupational concerns were common; these are stressors for many people during pandemics.[51] Pregnant women — and the public as a whole — would benefit from ensuring they are aware of financial assistance available during the pandemic and how and when it can be claimed.[54] In particular, the COVID-19 outbreak may be stressful for pregnant women who are 'critical workers' and therefore expected to continue working,[55] despite also being told they are a vulnerable group who should be 'particularly stringent in following social distancing measures'.[9] Organisations could help by changing the work roles of pregnant women, so they can work from home or away from the public where possible.

Participants in one study reported an overwhelming disruption in their expectations of birth, prenatal care and postnatal care, causing them to change their birth plans. In addition, maternity staff levels may be lower than usual during a pandemic

due to reassignment of staff to other areas of the hospital or staff minimising contact with patients for their own protection. This raises the question of what is an acceptable level of care to provide to uninfected pregnant women during a pandemic.[23] Guidance for healthcare professionals needs to be clear about which routine visits could be done over the phone or cancelled altogether, as well as how to provide appropriate care without exposing healthy women to illness. A solution may be designating a location and staff specifically for the care of healthy pregnant women.[56]

The literature showed that pregnant women often cope by taking drastic non-pharmaceutical precautions to avoid infection, which may affect all areas of their lives. They may become hypervigilant with regard to monitoring the most current self-protection information available, hygiene practices and reducing contact with others. These practices are recommended in infectious disease outbreaks and in themselves are positive behaviours as they reduce infection risk. However, it is possible that such measures could also cause distress. More research is needed to explore the benefits and risks to mental health of prolonged hypervigilance.

This review enhances understanding of how being pregnant during an emerging infectious disease outbreak may affect maternal mental health. Owing to the unpredictable nature of disease outbreaks, large numbers of women may find themselves pregnant during a pandemic, something they are unlikely to have expected or planned for. The psychological impact of pandemics may affect their mental health which could subsequently affect their children and families.

Overall, this review supports the suggestion that pregnant women are a highly vulnerable group in terms of psychological consequences during a pandemic;[57] they need to care for both their own health and that of their unborn child, in a 'doubling of health responsibilities'.[27] Planning for future pandemics should make considerations specific to pregnant women: involving them in pandemic preparedness exercises would ensure that their voices are heard and helping policymakers identify any gaps related to prenatal and postnatal care in current pandemic planning.

Limitations

Data screening, extraction and analysis were carried out by one author; in typical systematic reviews, it is preferable for double screening to take place and multiple reviewers to analyse the data. However, the results were discussed between all authors as the article went through multiple revisions before submission. Searches were limited to English language articles, meaning evidence may have been missed. No standardised quality appraisal of the included articles was carried out, as is common in rapid evidence reviews.[58] However, there were some particularly notable limitations to the literature, such as low response rates and a lack of quantitative research. Only one study[24] compared mental health outcomes for women pregnant during an outbreak with preoutbreak pregnant controls, making it difficult to ascertain the mental health–related differences in being pregnant during a disease outbreak and at any other time. No research directly compared pregnant women with non-pregnant individuals during an outbreak, so again, we cannot say whether pregnant women are more likely to experience stress during an outbreak than the general population. However, it is not unreasonable to think that the combination of usual pregnancy concerns and pandemic-related concerns may result in particularly negative psychological outcomes.

Conclusion

Pregnant women have specific needs during a pandemic and may be at risk of adverse psychological effects of the COVID-19 outbreak. This is important as there is a clear link between poor mental health in pregnant women and pregnancy complications. It is vital they are well informed about public health recommendations, which should include detailed description of benefits or lack of risk to unborn babies, as well as clear rationale for why prophylaxis or treatment is necessary. Virtual support groups specifically for pregnant women may be useful. Healthcare professionals involved in the care of pregnant women should be aware of the most current guidance and ensure that they closely monitor mental health during pregnancy and where necessary provide early evidence-based care.

Author statements

Ethical approval

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Competing interests

None declared.

Appendix I. Search strategy

1. pregnant OR pregnanc*
2. psychological OR mental health OR trauma OR stress OR distress OR anxiety OR well-being OR well-being OR panic OR depress*
3. pandemic* OR disease outbreak* OR SARS OR severe acute respiratory syndrome OR swine flu OR H1N1 OR avian influenza OR bird flu OR H5N1 OR Ebola OR MERS OR Middle East respiratory syndrome OR Zika OR coronavirus OR COVID-19.
4. 1 AND 2 AND 3.

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Short Communication

Risk of preventable injuries associated with Halloween

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ABSTRACT

Objectives: The objective of this study was to investigate the potential health risks associated with Halloween festivities.
Study design: This was a retrospective, population health study using insurance claims data between 2003-2014 representing more than 150 million unique Americans.
Methods: We analyzed the entire spectrum of external cause of injuries and quantified the relative risk associated with Halloween celebrations by comparing the observed diagnosis rate during Halloween week with its statistical expectation. We further used the closest federal holiday in October, Columbus Day, as a comparison to further corroborate the effects of Halloween.
Results: Our results indicate that no significant difference in relative risk for most conditions, like vehicle accidents, accidental poisoning and drowning, and adverse drug effects, during the Halloween season, when compared to the statistical expectation. However, we noticed a significant increase in the relative risk of accidental fall, self-inflicted injury, and injury inflicted by others, notably among young males.
Conclusion: Halloween is an exciting time of year for kids, families, and the entire community. A more vigilant approach toward celebration, including attempts to prevent fights and brawls, would help everyone have a safe and harmonious Halloween.

Introduction

Halloween is a secular holiday appreciated by children, adults, and a majority of US communities. Every year, millions of Americans participate in a variety of Halloween-related festivities. A recent study showed a 43% increase in all-age pedestrian fatalities on Halloween, with the death rate severalfold higher among children.[1] A handful of studies have investigated Halloween-associated hazards; however, the focus was primarily limited to traffic safety.[2,3] This study assessed putative, Halloween-associated health hazards within a spectrum of external injury causes, using the health insurance claims data set representing 150 million unique Americans.

Methods

To probe the risks associated with Halloween, we used the IBM Health MarketScan data set,[4] representing time-stamped health insurance claims between 2003 and mid-2014 for more than 150 million unique US individuals. We analyzed all claims related to external causes of injury and poisoning (as defined by the International Classification of Diseases version 9 codes E800–E999) and divided them into 23 subcategories. We then summarized each category's daily incidences for both biological sexes and in five age-groups (0–10, 11–20, 21–40, 41–60, and older than 60 years). We removed age-groups and sex groups with less than ten daily incident cases, resulting in a final set of 116 test groups, representing 15 external causes of injury and poisoning categories (Fig. 1).

Using a previously validated statistical model,[5] we quantified the relative risk (RR) associated with Halloween celebrations by comparing the observed diagnosis rate of a day around October 31 with its expectation, calculated by averaging diagnosis rates ± 2 weeks around the Halloween period. The diagnosis rate is the proportion of patients diagnosed with a condition of all enrollees at a specific point in time. Each year, Halloween falls on a different day of the week; therefore, some activities could continue over the weekends preceding or proceeding the Halloween day. To accommodate this fact, we studied seven days from October 28 to November 3 and also collapsed the day-level effects to obtain the overall week-level rates for Halloween week (Halloween day ± 3 days). Finally, the closest federal holiday in October, Columbus Day,

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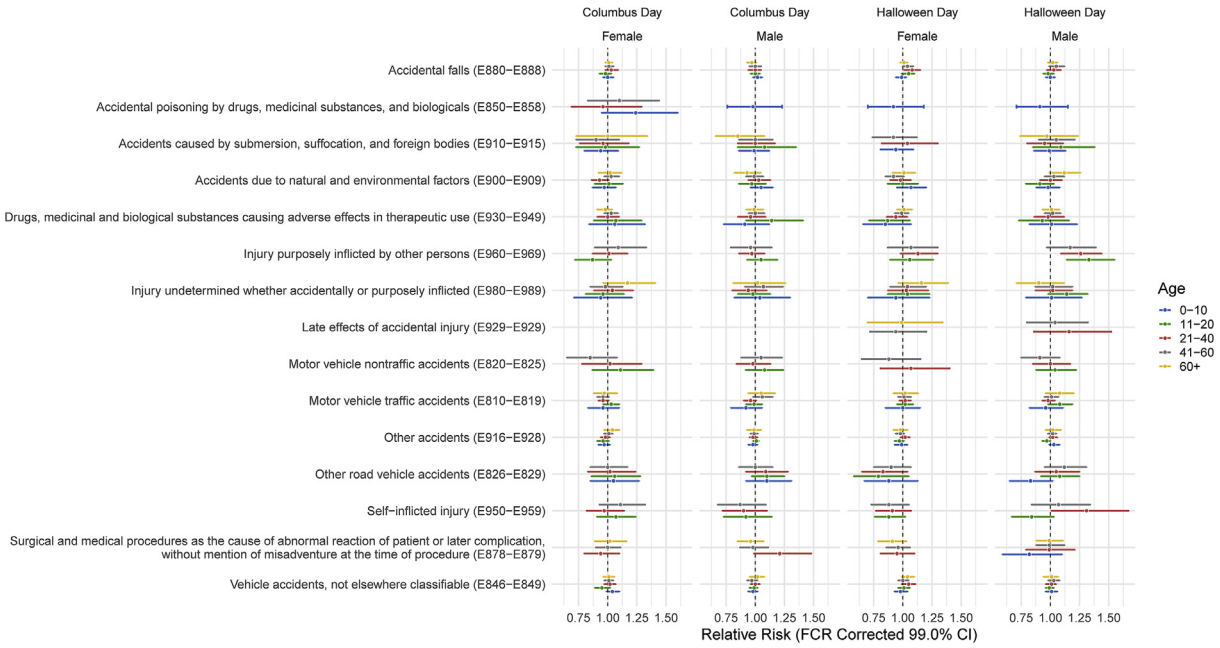

Fig. 1. Week-level relative risk estimates and corresponding 99.0% FCR-adjusted confidence intervals for the risks associated with Columbus Day and Halloween seasons. The RR estimates were computed for 15 categories (excluding 8 categories that have less than ten daily counts) and were stratified by sex and five different age-groups. There were no statistically significant variations in the risk of external cause of injuries around the Columbus Day period. The risk of accidental falls, self-inflicted injury, and the injury inflicted by others was elevated in some particular age-groups and sex groups during the Halloween period. FCR, false coverage rate; RR, relative risk; CI, confidence interval.

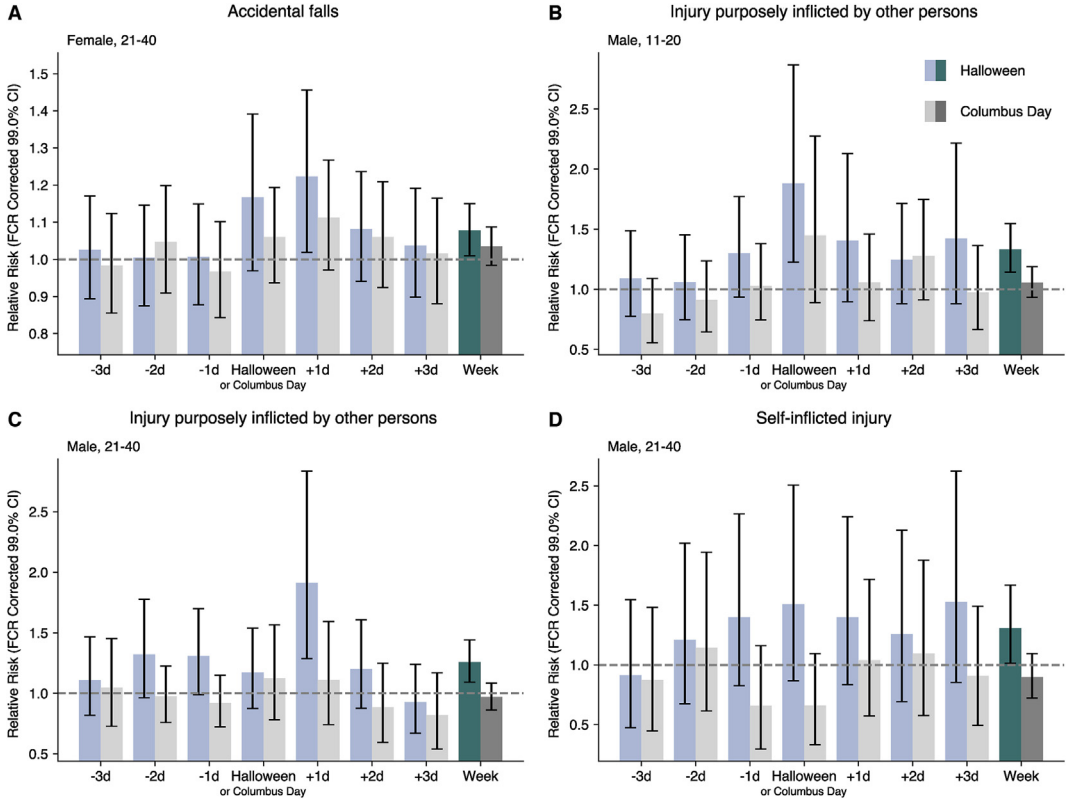

Fig. 2. The day-level relative risks (light blue bars: Halloween, light gray bars: Columbus Day) and corresponding 99.0% FCR-corrected confidence intervals for (A) homicides and injury inflicted by other persons among males aged 11 to 20 years; (B) homicides and injury inflicted by other persons among males aged 21 to 40 years; (C) suicide and self-inflicted injury among males aged 21 to 40 years; and (D) accidental falls among females aged 21 to 40 years. The individual bars represent day-level risk such as for Halloween, it ranges from October 28 (−3 d) to November 3 (+3 d). The rightmost green and dark gray bars against each subplot give the week-level effects, summarizing the entire-week risk from −3 days to +3 days. FCR, false coverage rate; CI, confidence interval. (For interpretation of the references to colour in this figure legend, the reader is referred to the Web version of this article.)

was used as a comparison to further corroborate the effects of Halloween.

We report week-level RR estimates for Halloween and Columbus Day along with the false coverage rate[6]–adjusted confidence intervals (CIs). The underlying aggregated study data, model specifications, and the day-level results are available on the project's website (https://github.com/hanxinzhang/halloween). The University of Chicago Institutional Review Board determined this study exempt from review, given that patient data were preexisting and de-identified.

Results

The results suggest no significant difference in RR for most conditions, such as vehicle accidents, accidental poisoning and drowning, and adverse drug effects, during the Halloween season, when compared with the expectation (Fig. 1). In addition, we found no significant change in risk for any of the studied categories around Columbus Day (Fig. 1).

However, we noticed a significant increase in the RR for four conditions around the Halloween period as compared with the statistical expectation: accidental falls in females aged 21–40 years (RR = 1.08, 99% CI = [1.01, 1.15]), injury inflicted by others in males aged 11 to 20 years (RR = 1.33, 99% CI = [1.14, 1.55]) and males aged 21 to 40 years (RR = 1.26, 99% CI = [1.09, 1.44]), and self-injury in males aged 21 to 40 years (RR = 1.31, 99% CI = [1.01, 1.67]). Fig. 2 shows the day-level and week-level RRs of those groups significantly elevated during the Halloween week, compared with the corresponding risk variation around Columbus Day. For the 11 to 20 age-group, risk peaked on Halloween day, whereas for the 21 to 40 age-group, risk mostly peaked the day after Halloween. Halloween's late-night activities (particularly those closer to midnight) perhaps delayed the appearance of symptoms from injuries, and less urgency to seek immediate medical attention among individuals aged 21 to 40 years may explain this distinctive trend.

Discussion

Based on our findings, and contrary to prior studies, we observe no significant increase in the risk of motor vehicle traffic accidents across all ages and sexes during Halloween. Instead, we found that the peril from Halloween celebrations is related to self-inflicted injuries and those inflicted by others, notably among young males. We believe these hazards are very much preventable and encourage efforts to minimize such risks. Halloween is an exciting time of year for kids, families, and the entire community. A more vigilant approach toward celebration, including attempts to prevent fights and brawls, would help everyone have a safe and harmonious Halloween.

Author statements

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Ethical approval

The University of Chicago Institutional Review Board determined this study exempt from review, given that patient data were preexisting and de-identified.

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Competing interests

The authors declare no competing interests.

Availability of data and material

The underlying aggregated study data, model specifications, and the day-level results are available on the project's website (https://github.com/hanxinzhang/halloween).

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Original Research

Socio-economic status and COVID-19–related cases and fatalities

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ABSTRACT

Objectives: The United States has the highest number of coronavirus disease 2019 (COVID-19) in the world, with high variability in cases and mortality between communities. We aimed to quantify the associations between socio-economic status and COVID-19–related cases and mortality in the U.S.
Study design: The study design includes nationwide COVID-19 data at the county level that were paired with the Distressed Communities Index (DCI) and its component metrics of socio-economic status.
Methods: Severely distressed communities were classified by DCI>75 for univariate analyses. Adjusted rate ratios were calculated for cases and fatalities per 100,000 persons using hierarchical linear mixed models.
Results: This cohort included 1,089,999 cases and 62,298 deaths in 3127 counties for a case fatality rate of 5.7%. Severely distressed counties had significantly fewer deaths from COVID-19 but higher number of deaths per 100,000 persons. In risk-adjusted analysis, the two socio-economic determinants of health with the strongest association with both higher cases per 100,000 persons and higher fatalities per 100,000 persons were the percentage of adults without a high school degree (cases: RR 1.10; fatalities: RR 1.08) and proportion of black residents (cases and fatalities: Relative risk(RR) 1.03). The percentage of the population aged older than 65 years was also highly predictive for fatalities per 100,000 persons (RR 1.07).
Conclusion: Lower education levels and greater percentages of black residents are strongly associated with higher rates of both COVID-19 cases and fatalities. Socio-economic factors should be considered when implementing public health interventions to ameliorate the disparities in the impact of COVID-19 on distressed communities.

Introduction

The coronavirus disease 2019 (COVID-19) pandemic is a worldwide public health crisis on a scale not witnessed in modern times. The United States has the highest number of COVID-19 cases and fatalities in the world.[1] There is wide variability in COVID-19–related mortality across countries and between communities within the U.S. While research has identified a number of risk factors for mortality such as age and comorbid disease, there are new data suggesting social determinants of health also influence outcomes.[2,3] Data from the Centers for Disease Control and Prevention suggest strong racial disparities in both COVID-19 prevalence and outcomes.[4] This point has been highlighted by the U.S. Surgeon General, the lay press, and recent publications.[4–6]

Socio-economic factors impact all facets of human functioning, including health-related quality of life. While a few publications have speculated on the impact of socio-economic status on the COVID-19 pandemic, there is limited quantitative analysis available.[7–10] A unique opportunity for investigation lies with the Distressed Communities Index (DCI). This metric was developed by the Economic Innovation Group and is a comprehensive estimate of socio-economic status by geographic location.[11] The DCI is a composite socio-economic ranking that accounts for unemployment, education level, poverty rate, median income, business growth, and housing vacancies. Previous studies have correlated a higher DCI score (lower socio-economic status) with worse health-related outcomes.[12–14] The objective of this study was to quantify the association between socio-economic status, both the composite metric of DCI and its individual component measures, and COVID-19 outcomes, accounting for other risk factors such as age and chronic diseases. We hypothesized that communities with lower

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socio-economic status (higher DCI scores) would have disproportionately higher COVID-19 prevalence and mortality.

Methods

Data sources and definitions

COVID-19–related cases and mortality were extracted from a publicly available data set of aggregated county sources (USAFacts) on May 2, 2020.[15] In addition, county-level estimates for chronic diseases were obtained from the Centers for Disease Control and Prevention. Specifically, the Centers for Disease Control and Prevention (CDC) COVID-19 response team used data from the 2018 Behavioral Risk Factor Surveillance System and U.S. Census population data to model prevalence of chronic diseases potentially associated with COVID-19 risk.[16] The model estimates for chronic kidney disease, chronic obstructive pulmonary disease, heart disease, obesity, and diabetes at the county level were included for risk adjustment. The data for rates of uninsured individuals under age 65, percent of population black, and percent of population age over 65 were obtained from the American Communities Survey 5-year Estimates. The United States is geographically and governmentally organized into 50 states. The states are then subdivided into counties with their own local governance structures. There is wide variability in the population and size of the states and counties, with the number of counties per state ranging from 3 (Delaware) to 254 (Texas). There is a total of 3143 counties or county-equivalent areas (parishes in Louisiana and boroughs in Alaska). At the time of writing, the most granular COVID-19–related data for the United States were available at the county level.

Socio-economic status was measured using the DCI and its components. A data use agreement is in place with the Economic Innovation Group.[11] The DCI score is available for 99% of the United States population and is derived from the American Communities Survey 5-year Estimates and Census Bureau County and Zip Code Business Patterns. The score is composed from the following data points: 'percentage adults with no high school degree, housing vacancy rate, percentage of adults not working, poverty rate, median income ratio, percentage change in employment, and percentage change in business establishments.' Detailed data definitions are available in Supplemental Table 1. The score ranges from 0 (no distress) to 100 (most distress) and is available on the zip code, county and state level. Severely distressed communities were defined by a DCI score >75. This study conformed to the principles embodied in the Declaration of Helsinki and was exempt from review by the University of Virginia Institutional Review Board due to the deidentified nature of the publicly available data sets.

Counties were included for analysis if both COVID-19 data and DCI score were available. Individual cases and deaths were excluded from the analysis if they were not attributed to a specific county.

Statistical analysis

Categorical variables are presented as counts (%) and continuous variables as median [25th, 75th percentile]. Counties were stratified by severely distressed (DCI>75) versus less distressed (DCI $\leq$75) for univariate analysis using the Mann-Whitney U test. Adjusted effects of socio-economic status were approximated using hierarchical linear mixed models with Laplace approximation and a negative binomial distribution. Socio-economic effects were adjusted for county-level prevalence of elderly residents, uninsured adults, and the following comorbidities: chronic kidney disease, chronic obstructive pulmonary disease, heart disease, diabetes, and obesity. A random effect was used for clustering at the state level. Temporal bias was accounted for with a variable of days since first COVID-19 case. Regression results were exponentiated to construct adjusted rate ratios. Statistical analyses were carried out using SAS, version 9.4 (SAS Institute, Cary, NC), with some graphical representation performed with Prism 8 (GraphPad, San Diego, CA). A *P*-value <0.05 determined statistical significance.

Results

Cohort description

At the time of data extraction on May 2, 2020, there were 1,096,706 cases and 64,066 deaths in the United States. There were 3143 counties with COVID-19 data available. Twenty-four states had cases not allocated to counties for a total of 6465 cases (0.6%), while 14 states had unallocated deaths for a total of 1754 deaths (2.7%), with 1445 being unallocated in Tennessee. A DCI score was available for 3127 counties with COVID-19 data. This cohort included 1,089,999 cases (99%) and 62,298 deaths (97%) for a case fatality rate of 5.7%. A consort diagram depicting the exclusions leading to the final cohort is shown in Fig. 1.

Differences by the DCI

A total of 781 counties were designated as severely distressed (DCI >75), accounting for 68,741 (6.3%) COVID-19 cases and 3811 (6.1%) COVID-19 fatalities. When compared with less distressed counties, those classified as severely distressed had significantly lower median COVID-19 cases (13 [3–39] vs 21 [4–117], $P < 0.0001$) and fatalities (0 [0–1] vs 0 [0–4], $P < 0.0001$). Owing to the rural nature of many distressed counties, after accounting for county population size, there was no difference in median cases per 100,000 persons and a reverse in the trend with higher median fatalities per 100,000 persons in severely distressed counties (Table 1). The median percentage of black Americans was significantly higher in severely distressed counties compared with less distressed counties (7.3% vs 1.8%, $P < 0.0001$). Other significant differences included severely distressed counties having higher rates of elderly residents, uninsured individuals, and individuals with chronic kidney disease, chronic obstructive pulmonary disease, heart disease, diabetes, and obesity. However, severely distressed communities were hit later in the pandemic with lower median days since first case (34 vs 38, $P < 0.0001$).

Risk adjusted analyses of socio-economic status

Cases of COVID-19 by county were associated with multiple socio-economic factors and days since first infection but not health-related comorbidities. The adjusted risk ratios can be visualized in Fig. 2, with complete model results in Table 2. A higher number of cases were associated with lower education level, higher proportion of black Americans, higher income, and lower poverty rate. The specific covariates with significant associations with cases per 100,000 persons were the percentage of adults without a high school degree (RR 1.10), proportion of black residents (RR 1.03), median income ratio (RR 1.01), and poverty rate (RR 0.98). Fig. 3 shows risk-adjusted predicted cases per 100,000 persons for the two strongest associations, the percentage of adults without a high school degree and proportion of black residents. No comorbid medical conditions were associated with COVID-19 cases at the county level.

Higher COVID-19 mortality was associated with higher income but lower education, higher employment rate, higher proportion of black Americans, older residents, and less obesity (Table 2). The significant socio-economic associations were the percentage of adults without a high school degree (RR 1.08), median income ratio

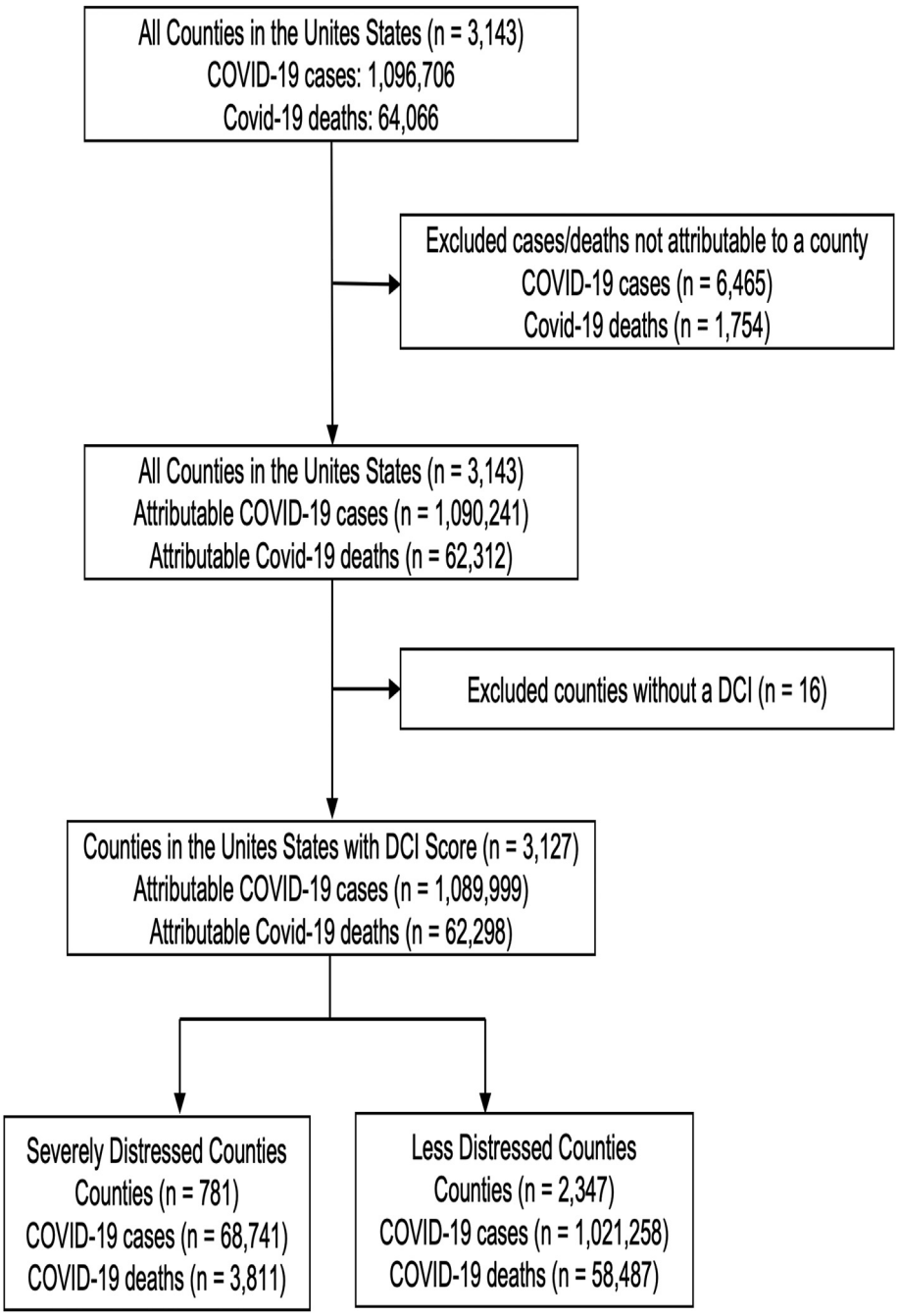

Fig. 1. Consort diagram of included and excluded counties, COVID-19 cases, and COVID-19 fatalities. COVID-19 = coronavirus disease 2019.

(RR 1.01), and percentage of adults not working (RR 0.98). Significant demographic and health-related associations included the percentage of population aged older than 65 years (RR 1.07), proportion of black residents (RR 1.03), and obesity prevalence (RR 0.96). Fig. 4 shows the risk-adjusted predicted fatalities per 100,000 persons for the two strongest socio-economic associations, the percentage of adults without a high school degree and proportion of black residents.

Discussion

The present study highlights community-level socio-economic disparities in COVID-19 prevalence and mortality within the United States. Severely distressed communities account for a disproportionately higher number of deaths per 100,000 persons. However, poverty and unemployment were protective against contracting COVID-19, highlighting the unique risks of a pandemic and limitations of composite metrics of socio-economic status. Risk-adjusted analysis identified two socio-economic determinants of health as having the strongest association with both population-adjusted cases and fatalities: lower education level and higher proportion of black residents. Although age, comorbidities, and overall health status also appear to impact COVID-19 prevalence and outcomes, socio-economic factors have higher risk ratios and represent the strongest associated factors with both number of cases and fatalities.

Table 1
Univariate analysis by severely versus less distressed counties.

Variable	Severely distressed	Less distressed	*P*-value
Cases per 100,000 persons	61 [22–176]	63 [27–145]	0.813
Deaths per 100,000 persons	0 [0–7.1]	0 [0–5.4]	0.028
% of population aged older than 65 years	17.4% [15.4–19.5]	16.8% [14.1–19.6]	<0.0001
% of population Black	7.3% [1.1–33.1]	1.8% [0.6–6.7]	<0.0001
% of uninsured population aged younger than 65 years	14.9% [11.1–17.5]	10.0% [7.0–14.0]	<0.0001
Chronic kidney disease prevalence	4.0% [3.7–4.3]	3.3% [3.0–3.5]	<0.0001
COPD prevalence	11.0% [9.8–12.3]	8.3% [7.0–9.8]	<0.0001
Heart disease prevalence	10.1% [9.3–11.0]	8.2% [7.1–9.1]	<0.0001
Diabetes prevalence	15.8% [14.5–17.3]	12.0% [10.8–13.5]	<0.0001
Obesity prevalence	38.1% [35.7–40.5]	34.6% [31.7–36.9]	<0.0001
Days since first case	34 [25–39]	38 [29–44]	<0.0001

All values expressed as median [interquartile range]; COPD = chronic obstructive pulmonary disease.

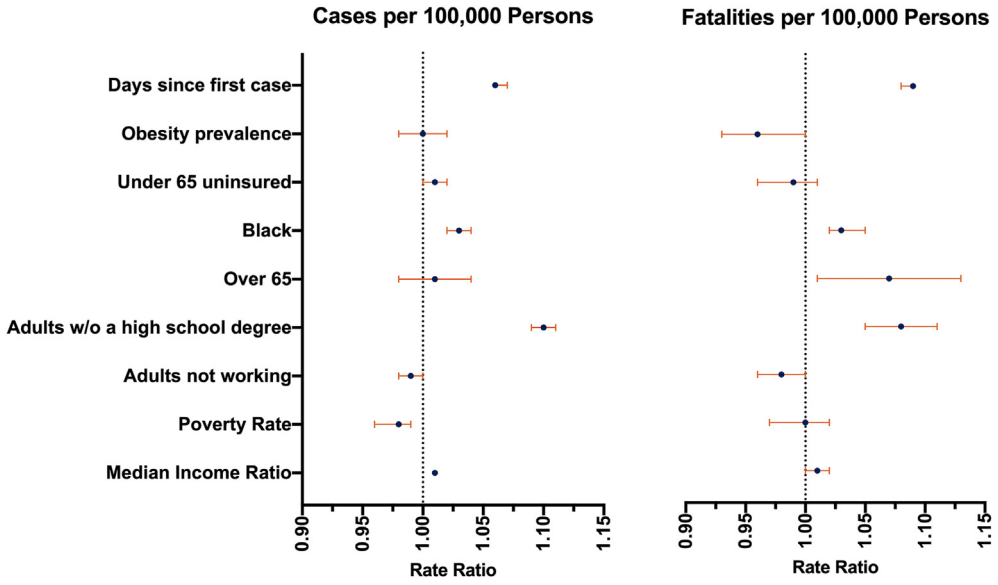

Fig. 2. Forrest plot of adjusted rate ratio for covariates significantly associated with either COVID-19 cases or fatalities per 100,000 persons. COVID-19 = coronavirus disease 2019.

Table 2
Adjusted rate ratios for COVID-19–related cases and fatalities per 100,000 persons.

Variable	Cases per 100,000 persons			Fatalities per 100,000 persons		
	Rate ratio	95% CI	*P*-value	Rate ratio	95% CI	*P*-value
Housing vacancy rate	1.01	1.00–1.02	0.1382	0.98	0.95–1.01	0.1274
Median income ratio	1.01	1.01–1.01	<0.0001	1.01	1.00–1.02	0.0094
Poverty rate	0.98	0.96–0.99	0.0019	1.00	0.97–1.02	0.8511
% change in employment	1.00	1.00–1.00	0.2358	0.99	0.99–1.00	0.0521
% change in establishments	1.00	0.99–1.01	0.7611	1.00	0.99–1.02	0.7516
% of adults not working	0.99	0.98–1.00	0.1496	0.98	0.96–1.00	0.0266
% of adults w/o a high school degree	1.10	1.09–1.11	<0.0001	1.08	1.05–1.11	<0.0001
% population in distressed zip codes	1.00	1.00–1.00	0.7609	1.00	0.99–1.00	0.2284
% of population aged older than 65 years	1.01	0.98–1.04	0.410	1.07	1.01–1.13	0.0216
% of population Black	1.03	1.02–1.04	<0.0001	1.03	1.02–1.05	<0.0001
% of uninsured population aged younger than 65 years	1.01	1.00–1.02	0.1665	0.99	0.96–1.01	0.2239
Chronic kidney disease prevalence	0.81	0.47–1.37	0.4266	0.84	0.34–2.09	0.7136
COPD prevalence	1.00	0.89–1.13	0.967	1.15	0.95–1.41	0.1576
Heart disease prevalence	1.03	0.81–1.32	0.7833	0.89	0.58–1.37	0.5919
Obesity prevalence	1.00	0.98–1.02	0.7464	0.96	0.93–1.00	0.0358
Diabetes prevalence	0.98	0.89–1.09	0.7132	1.06	0.89–1.26	0.5221
Days since first case	1.06	1.06–1.07	<0.0001	1.09	1.08–1.09	<0.0001

CI = confidence interval; COPD = chronic obstructive pulmonary disease; COVID-19 = coronavirus disease 2019.

Using the composite DCI metric, which includes seven different socio-economic factors, we found that counties with lower socio-economic status (higher DCI score) have higher COVID-19 death rates per 100,000 persons compared with non-distressed counties. Rates of high mortality in distressed communities are multifactorial, likely accounting for differences in age, number of chronic medical conditions per person, and socio-economic status. One of the strongest predictors of mortality from COVID-19 is age, where

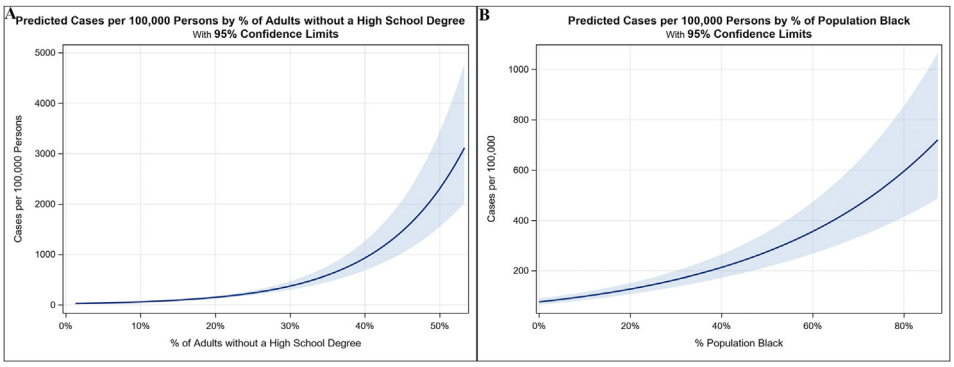

Fig. 3. Adjusted predicted COVID-19 cases per 100,000 persons for the two strongest socio-economic factors. (A) Percentage of adults without a high school degree and (B) percentage of population Black. Estimates are shown holding all other predictors at mean values (Supplemental Table 1). COVID-19 = coronavirus disease 2019.

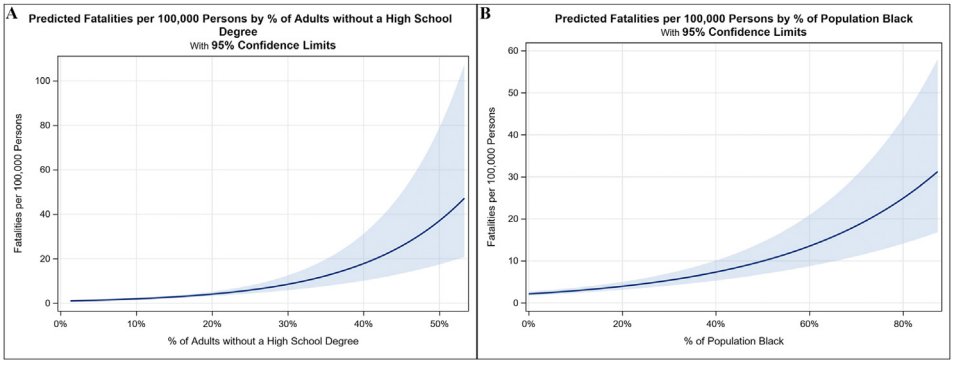

Fig. 4. Adjusted predicted COVID-19 fatalities per 100,000 persons for the three strongest socio-economic factors. (A) Percentage of adults without a high school degree and (B) percentage of population Black. Estimates are shown holding all other predictors at mean values (Supplemental Table 1). COVID-19 = coronavirus disease 2019.

in Italy the case fatality rate in octogenarians was more than 20%.[3,17] Distressed communities are rapidly aging and have a greater proportion of elderly residents.[18] In addition, lower socio-economic status has been associated with a higher burden of chronic medical conditions and comorbidities.[13] These underlying medical comorbidities have also been shown to increase COVID-19–related mortality.[2] However, in the present data, we demonstrate a stronger association between socio-economic factors than underlying medical conditions with only obesity inversely associated with COVID-19 mortality. Obesity has traditionally been considered protective against acute respiratory distress syndrome–related mortality, but preliminary data on COVID-19 suggest possible increased risk.[19,20] There are significant overlaps in socio-economic status and obesity that were not accounted for in other studies, and the independent effect of obesity of COVID-19–related cases and mortality requires further study.

Distressed communities tend to be disproportionately comprising black Americans, and these individuals have higher COVID-19 mortality rates.[4–6,21] A study not yet published found that communities with a higher proportion of black residents had increased COVID-19 cases (relative risk 1.24) and deaths (relative risk 1.18).[22] Our study corroborates these findings where a higher proportion of black residents were associated with both increased cases and fatalities per 100,000 persons. The exact reasons to explain these associations cannot be determined from the current data sets and analysis, and a more thorough understanding of racial inequities requires nationwide disaggregated data. The underlying etiology for the racial inequity likely reflects trends of economic, geographic, and health disparities seen with socioeconomically disadvantaged populations. In addition, the impact of systemic flaws in the structural organization of American society with racism, access to health care, and community-level resources should not be underestimated.

The COVID-19 pandemic in America is becoming increasingly a story of economics. Significant attention has focused on how public health measures to save lives (social distancing, closure of all non-essential business, stopping the delivery of all non-urgent/emergent health care) have resulted in one of the largest increases in unemployment in American history. However, our data focus on the opposite relationship, how economics is affecting who contracts and dies from COVID-19.[23] Those individuals in frontline jobs who maintain employment are at increased risk of contracting coronavirus, particularly with concerns related to asymptomatic spread in the setting of inadequate testing.[24] Our data are supportive of this theory in two ways. First, higher income and less poverty both were associated with higher rates of COVID-19 cases while higher unemployment was associated with fewer COVID-19 fatalities. These findings suggesting a significant amount of transmission is through the workplace. Second, lower education levels were also associated with contracting COVID-19. This could be due to frontline jobs being disproportionately low paying and without a requirement of an advanced degree.[23] The education relationship is relatively strong (RR 1.10) compared with other metrics, which may be due to the highest paying jobs being more amenable to remote working, making employment and income imperfect markers of risk.

This public health trope of economics underpinning health inequalities is true not only in America but also was highlighted in 1980 in Britain with the Black Report.[25] One of the expectations of establishing the National Health Service with universal healthcare

coverage was to eliminate health disparities. However, thirty years after, it was made clear that economic divisions (employment, income, education, housing, and so on) were only widening health disparities in the United Kingdom. The next few decades saw focused efforts to identify the causes of morbidity and mortality differences, develop better metrics to gauge true health disparities, and quantify those differences.[26–28] COVID-19 represents a rare disruptive opportunity to refocus our efforts on using this knowledge to reverse health disparities.

Now is the time to be proactive in the management of this pandemic and focus allocation of federal and state-level resources to these distressed communities. As society attempts to safely reverse restrictive public health measures such as stay at home orders, other more resource-intensive tools including testing and contact tracing will be required. The DCI may help to guide public health interventions to areas most in need. Clear and decisive actions are needed to help safeguard vulnerable populations who are at the highest risk for complications and death.[29]

This study is limited by the lack of epidemiologic data available at the zip code level or widespread detailed demographic information, representing systemic failures of the healthcare system which limits our ability to adequately respond in an equitable manner. Socio-economic factors are interrelated, and collinearity limits the utility of an epidemiologic study design such as this. Individual-level SES analyses are required to better understand the interaction of all aspects of Socioeconomic status on COVID-19. In addition, the delay in testing availability and narrow testing criteria vastly overestimate the case fatality rate, but it is unclear how this varies across the United States. Better racial data are needed regarding COVID-19 to truly understand the disparities by race beyond a rough estimation for black Americans. Finally, the DCI only accounts for a small portion of variability in case fatality rates, suggesting social determinants of health are only one factor in determining COVID-19–related outcomes.

In summary, socio-economic determinants of health are associated with COVID-19 prevalence and mortality. Severely distressed counties with low socio-economic status have higher rates of both COVID-19 cases and fatalities than communities with higher socio-economic status. The socio-economic influence is broad, with many components of the DCI (education level, income, and poverty rates) being associated with COVID-19 cases. However, due to the transmission patterns of COVID-19 in the United States, higher income but lower education levels were associated with COVID-19 cases and fatalities. A higher proportion of black residents were also one of the strongest associations with COVID-19 cases and fatalities. Socio-economic factors and the DCI could help allocate appropriate public health resources in areas with lower socio-economic status, a critical step toward an equitable fight against COVID-19.

Author statements

Ethical approval

This study is exempt from IRB review.

Funding

No funding was utilized for this research.

Competing interests

The authors report no conflicts of interest.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2020.09.016.

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Short Communication

Spatial relationship between ambulatory surgery centers and colorectal cancer mortality within Pennsylvania, United States

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ABSTRACT

Objectives: The objective of this study was to evaluate the spatial relationship between colorectal cancer (CRC) mortality and ambulatory surgery center (ASC) density in Pennsylvania's 67 counties.
Study design: This was an ecological study.
Methods: Age-adjusted CRC mortality rates were linked to ASC densities per 1,000 people. The data set was analyzed using global, local, and regional Moran's I, to test for randomness in CRC mortality and ASC density.
Results: CRC mortality rates (median: 15.30 per 100,000 of the US 2000 standard million population) exhibited hot spots in rural Pennsylvania counties. ASC densities (median: 0.35 providers/km^2 per 1,000 people) showed hot spots in urban southeastern Pennsylvania and cold spots in northern Pennsylvania.
Conclusions: CRC mortality rates tended to cluster in rural northern Pennsylvania counties; ASC density tended to cluster in urban southeastern counties, indicating a spatial disparity between needed and provided healthcare resources. There is a need for public health and health system changes to increase the availability of CRC services to rural communities.

Colorectal cancer (CRC) is the second leading cause of cancer death in the United States and in Pennsylvania, United States.[1] Spatial variation in CRC treatment and mortality has been reported; few studies have examined the spatial relationship between CRC screenings, such as endoscopies, which have been shown in studies to reduce CRC mortality.[2–4] Screening endoscopies are typically performed at an ambulatory surgery center (ASC), an outpatient facility that performed about one million endoscopies in 2018.[5] The objective of this study was to understand and to evaluate the spatial relationship between CRC mortality and ASC density in Pennsylvania's 67 counties.

We extracted 2013–2017 county-specific, age-adjusted CRC mortality rates (ICD-10: C18–C20 and C260) from the Enterprise Data Dissemination Informatics Exchange. These data were collected and maintained by the Pennsylvania Department of Health. The dependent variable was the age-adjusted CRC mortality rates in 67 counties (median: 15.30 per 100,000 of the US 2000 standard million population; Shapiro-Wilk normality: $P < 0.001$; skewness: −1.98), using SPSS 26.

The ASC locations were obtained from the December 2019 Provider of Services file compiled by the Centers for Medicare & Medicaid Services. In 2019, there were 149,207 total providers, 5,696 of which were in Pennsylvania. There were 391 ASCs in the state, with 106 offering colonoscopies or endoscopies. The 106 ASC point estimates were then aggregated to the county using zonal statistics for the density numerator. County shape files, with areal measurements for land and water area (m^2), were obtained from the United States Census Bureau's TIGER/Line Shapefiles. For the density denominator, we used total area, which was converted to square kilometer (International System of Units [SI]: 100,000 m^2) for easier interpretation. For the independent variable estimation, we used the ASC point estimates and areal units in square kilometer to calculate the density of facilities per 1,000 people in 67 counties (median: 0.35 providers/km^2 per 1,000 people; Shapiro-Wilk normality: $P < 0.001$; skewness: 5.03), using SPSS 26.

For the spatial analysis, we projected the data using the World Geodetic System 1984 coordinate system. We used global Moran's I in GeoDa 1.14 to test for randomness of county-specific CRC

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mortality rates and for ASC density. For the definition of the spatial analysis matrix, we created an ID variable and used queen contiguity weights with an order of contiguity of one and no precision threshold. This choice was justified because it deals with potential inaccuracies in the polygon file, such as rounding errors. We performed univariate global and local indicator of spatial association (LISA; 99,999 permutations) Moran's I clustered analyses to examine the spatial autocorrelation between CRC mortality and ASC density. Edge effects were corrected for by brushing the spatial weights, which corrected for locations that were not part of the selection or unselected observations, creating a regional Moran I autocorrelation in univariate global and LISA estimates. Regime regression coefficients were provided for both hot spots and cold spots, with at least two neighbors (i.e. selected) vs. unselected counties. Counties with positive z-scores and $P < 0.05$ indicate a hot spot; counties with negative z-scores and $P < 0.05$ indicate a cold spot. The high-high and low-low locations (positive local spatial autocorrelation) are typically referred to as spatial clusters, whereas the high-low and low-high locations (negative local spatial autocorrelation) are termed spatial outliers. The higher (or lower) the z-score, the more intense the clustering; z-scores near zero means no spatial clustering. Note that the LISA estimates are a relative measure and can only be interpreted within the context of its computed z-score or *P*-value.

The spatial analyses results showed that CRC mortality rates had a marginally significant (P < 0.10) univariate global spatial autocorrelation (Moran's I [r]: -0.12, z-score [z]: -1.46, pseudo P value [p]: 0.056). Additionally when doing an univariate LISA estimation, CRC mortality rates in Pennsylvania had statistically significant spatial hot spots in the three southwest counties (selected: [r:-0.50; z: <-20, p: <0.001]; unselected: [r:-0.17, z: -2.53, p: <0.001]) and either low-high or low-high (selected: [r: -1.01; z: 0.14, p: 0.336]; unselected: [r: 0.12, z: 3.17, p: 0.001]) in the four rural north central Pennsylvania counties (Fig. 1). ASC densities per county were found to have a statistically significant (P < 0.05) and positive univariate global spatial autocorrelation (r: 0.46; z: 8.51; p: <0.001). Hot spots for ASC density were observed in four southeast counties (selected: [r: -0.45, z: -7.38, p: <0.001]; unselected: [r:-0.17, z: 14.52, p: <0.001]). Cold spots were observed in eight northern counties (selected: [r: 0.13, z: -0.032, p: 0.50]; unselected: [r: 0.45, z: 6.65, p: <0.001]). Lastly, the Spearman's correlation between CRC mortality and ASC density was weak (-0.22).

In our study, CRC mortality tended to cluster in rural northern Pennsylvania counties. These findings are consistent with other studies that identified barriers (i.e., physician shortages in rural counties) that may prevent people from engaging in preventive care in rural areas.[6,7] Another study examined overall trends and concluded that CRC rates are declining, but their sample was limited to US adults aged 50 years or older and did not take into consideration rurality or rates for younger age-groups.[8] Although national surveillance data suggest that CRC mortality is trending downward, further research may be necessary to explore the relationship between CRC mortality and access to care to ensure that these decreases are equitably distributed.

ASCs perform approximately 22.5 million procedures per year, specifically performing 4.0 million colonoscopies and 2.2 million endoscopies of small intestines in 2010; only 2% of ASC visits had discharges to an inpatient hospital.[5] ASCs performing endoscopies or colonoscopies were highly clustered in urban southeastern Pennsylvania and less clustered in rural northern counties. This was one of the few studies to use provider density (km^2) rates per 1,000 people, which was adapted from the World Health Organization[9] and from a mammogram facility study.[10] Our study, as well as studies of mammography, found spatial disparities in clustering of ASCs, based on densities per 1,000 people.

This study is not without limitations, including the ecological fallacy, skewness, edge effects, and the modifiable areal unit problem. The ecological fallacy arises when using an area-based measure (e.g., ASC density) as a surrogate for individual-level characteristics (e.g., individuals receiving a colonoscopy at an ASC). ASC density and CRC mortality were both highly skewed in opposite directions (i.e., mortality: left skewed; ASC density: right skewed), meaning that z-scores in the LISA are not normally distributed, but this issue was mitigated by using 99,999 permutations. In addition, there may have been an edge effect because the mortality data were limited to Pennsylvania. However, within Pennsylvania, the edge effect was partially mitigated statistically by using regional Moran's I analyses. The modifiable areal unit problem arises when a geographic area (e.g., county) is defined by a boundary that is created from non-related criteria (e.g., geopolitical factors) or aggregation bias. However, because the data used a larger area and not census block or below and Spearman's correlation coefficients were close to ±0.30, i.e., a weak effect, the modifiable areal unit problem was mitigated. This study used an ecological approach versus point-based estimation because of data availability and privacy concerns. The mortality rates were only available as age-adjusted rates at the county level.

Despite these limitations, this analysis has notable strengths. It is an ecological study, which is population based and provides an exploratory look at the relationship with ASCs and CRC mortality. An individual-level study is not possible, given that the CRC mortality rates are provided as ecological variables, for which there is no correlate at the individual level. The study was the first to examine CRC facility densities using data from ambulatory service centers providing endoscopies in Pennsylvania. Finally, it expands the spatial work on CRC mortality including ASC facility density.

In conclusion, the results help explain spatial variation in CRC mortality and identify locations for future public health interventions. This analysis will help inform prevention and screening programs in Pennsylvania's counties where there are unusually high rates of CRC mortalities. One recommended

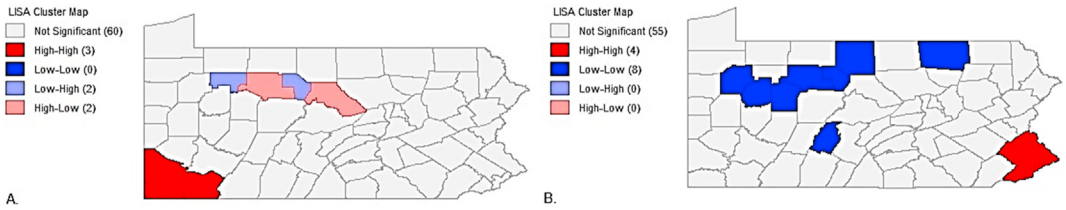

Fig. 1. Univariate local indicator of spatial association (LISA) autocorrelation cluster maps for (A) age-adjusted colorectal cancer mortality rate (per 100,000 of the US 2000 standard million population) and (B) ambulatory service center density (providers/km^2 per 1000 people) within 67 Pennsylvania counties.

change is offering remote or home-based medical services to rural areas to overcome the challenge of lower access to ASCs. There is an ongoing need for systematic spatial approaches to analyzing CRC screening and treatment providers relative to CRC mortality rates. Finally, there is a need for potential policy and health system changes to increase the availability of CRC services to rural communities.

Author statements

Author Contributions

Conceptualization, E. J. Lengerich and N. R. Geyer; methodology, N. R. Geyer and M. Wang; software, N. R. Geyer; formal analysis, N. R. Geyer; resources, E. J. Lengerich; data curation, N. R. Geyer; writing—original draft preparation, N. R. Geyer; writing—review and editing, J.L. Moss and M. Wang visualization, N. R. Geyer; supervision, E. J. Lengerich; project administration, E. J. Lengerich; funding acquisition, E. J. Lengerich. All authors have read and agreed to the published version of the manuscript

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Ethical approval

This study was approved by the Penn State College of Medicine Institutional Review Board (ID = STUDY00003894).

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Competing interests

None declared.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2020.09.018.

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Original Research

The lagged effect of state gun laws on the reduction of state-level firearm homicide mortality in the United States from 1999 to 2017

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ABSTRACT

Objectives: Gun violence is a pressing concern in the United States, where many laws targeting gun violence vary across states and localities. Studies have investigated the association between gun laws and gun violence, but not many focus on the role of time, which is critical for implementation. This study aims to determine the lagged association of gun laws with firearm homicide mortality to better understand the impacts of state gun laws over time.
Study design: The design of this study is a longitudinal study.
Methods: This study applied the distributed lag non-linear model to assess the lagged association between firearm homicide mortality and the number of gun law provisions at the state level from 1999 to 2017. State gun law provisions were analyzed in aggregate and also across five groups of regulations. All estimates were transformed into relative risks (RRs).
Results: Regarding all state gun laws, regardless of how many gun law provisions were on the books in any year, a significantly reduced RR of firearm homicide mortality was not observed until 7 years later. Among the five regulation groups, a significant RR less than 1 was more likely to happen in longer lags $\geq$5. The lowest significant RR = 0.24 (95% confidence interval = 0.15, 0.39) was observed in the gun types, components, and trafficking group at lag 8. All regulation groups had an overall impact to reduce significantly the RR of firearm homicide mortality with more gun law provisions, except for the dealer regulation group.
Conclusions: State gun law provisions appear to impact firearm homicide mortality differently over time. This emphasizes the centrality of enforcement. Firearm policy researchers need to consider how specific gun laws are implemented over time to help inform law-based interventions.

Introduction

Gun violence is as a major public health issue in the U.S. High profile mass shootings can take the lives of multiple people in one incident. But gun violence is also a problem on days without mass shootings. Firearm homicides are the fourth leading cause of injury deaths among children aged 5–14 years, the third leading cause of injury deaths among youth aged 15–24 years, and the fifth leading cause of injury deaths among all ages in the U.S.[1] Victims of firearm-related homicides are disproportionately men, African American, and in urban areas.[2] Between 1999 and 2017, cumulative firearm-related deaths were responsible for 6.8% of all years of potential life lost.[3] Despite these numbers, gun regulation in the U.S. is controversial.[4] Gun regulation is also laden with political barriers, including engaged gun owners but gun safety advocates who, at least historically, 'have struggled to develop an organizational infrastructure, stable funding sources, savvy legislative strategies, and broadly resonant messages to match those of their opponents.'[5]

The regulation of guns in the U.S. involves an interaction of federal, state, and local government actions. For example, the federal Brady Handgun Violence Prevention Act requires federally licensed firearm dealers to conduct background checks of potential buyers, yet many guns are acquired from those who are not federally licensed firearm dealers.[6] States and localities too have enacted and implemented laws to regulate gun transactions, storage, and possession, including in the area of background checks.[7]

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All of these federal, state, and local governmental actions operate with a backdrop of jurisprudence interpreting the Second Amendment of the U.S. Constitution, which the U.S. Supreme Court has recently interpreted to protect a limited right of individuals to possess guns in their homes for self-defense.[8,9] However, the court simultaneously reinforced the validity of many 'longstanding prohibitions,' such as banning carrying firearms in places such as schools and prohibiting firearm possession by felons.[8]

Individual studies about which gun laws reduce gun violence, which are necessary for effective regulation, have provided important insights. For instance, at the federal level, restriction on firearm access by those convicted of misdemeanor domestic violence offenses is associated with a reduced relative risk (RR) of intimate partner homicide using a firearm,[10] while implementation of the Brady Act has prevented millions of gun sales to disqualified individuals even though most weapons used in crimes are not obtained from covered dealers.[11] The meta-research also suggests that certain gun laws may help move the needle on gun violence. Some gun regulations, particularly background check laws and laws requiring buyers to obtain permits before purchase, appear linked to decreased homicide rates.[7] Current evidence is moderate to conclude that background checks, prohibitions associated with mental illness, and stand-your-ground laws decrease firearm homicides or violent crime and only limited to conclude that concealed carry laws increase violent crimes. However, it is inconclusive on whether many other types of gun policies move the needle on violent crime, as well as suicide and unintentional injuries and deaths.[12] Despite the helpful research about the effectiveness of gun law impact on violence, unanswered questions also remain.[7,13–15]

In the context of gun violence, law can function as a real or perceived obstacle to gun violence interventions or as an intervention itself.[16] The model to help understand law as a gun violence intervention posits that adopted laws – as implemented – can modify behaviors, environments, or social norms (or all of the aforementioned) to impact health outcomes. However, all of this can take time and involve implementation through different practices.[17] Therefore, determining when gun-related injury rates meaningfully change in the presence of gun-related laws is important because it can help identify what time periods to further examine to unearth legal practices, behaviors, environments, and social norms that may be contributing to the changes in gun-related violence stemming from gun-related laws and their enforcement.

This study aims to research the longitudinal association between state gun law provisions on the books and firearm homicide mortality. We hypothesized that the reduction of firearm homicide mortality would not be immediate but would occur over time. We analyzed yearly firearm homicide death counts and existing state-level gun law provisions across the 50 states from 1999 to 2017. We expected to help identify the gap between the changes to state-level gun laws and the reduction of firearm homicide mortality so that these identified time periods can be studied further.

Methods

Data sources

We queried the Centers for Disease Control and Prevention Wide-Ranging Online Data for Epidemiologic Research by state and year from 1999 to 2017 to fetch annual homicide deaths from firearms, which included all ages, genders, races, and ethnic origins.[18] For a few states with suppressed data (i.e., under 10 cases), an advanced data query was used. This query involved using data from one state with a suppressed value and another state without a suppressed value and deriving the suppressed data from the combined data through subtraction.

Gun law data were obtained from the State Firearm Law Database, which catalogs 134 state-level gun law variables by year across 50 states.[19] This longitudinal catalog is constructed to facilitate evaluations of state-level gun laws and focuses on key firearm safety provisions.[20] The database classified gun law provisions into several categories. We first counted the total number of gun law provisions in place in each state every year, where more provisions appear to indicate greater state-level restrictions. Next, we merged these categories into five broader groups in terms of the buyer regulation group, the dealer regulation group, the domestic violence regulation group, the gun types, components, and trafficking regulation group, and the possession, assess, and carrying regulation group. Details about the five regulation groups appear in Table 1.

We also considered confounding variables often used in gun violence literature. These were mainly socio-economic status variables captured in the American Community Survey,[21] including the percentages of residents who are black, men aged 15–34 years, unemployed, divorced, and of households in poverty. We also considered violent crime and law enforcement officer rates obtained from the U.S. Federal Bureau of Investigation Uniform Crime Reports.[22]

Statistical modeling approach

To verify the non-linearity between the gun law and firearm mortality variables, we initially adopted the generalized additive model to fit the number of the firearm mortality by the spline of number of total gun laws and specific regulation category in each lag, respectively. The results show that all splines are less than 0.0001. Thus, we further applied the distributed lag nonlinear model to investigate the association between the presence of gun law provisions and gun violence over time.[23,24] Thus, assume that Y_{it} is the number of firearm homicide death at time t and state i, which follows a Poisson distribution with a mean parameter μ_{it}. Thus, the model based was built as follows:

$$\begin{aligned}\mathrm{Log}(\mu_{it}) = {} & \alpha + \alpha_i + f(\mathrm{LAW}_{it}, L = k) \\ & + \boldsymbol{\beta} \times (\mathrm{Confounders})_{it} + f(t) + \mathrm{offset}_{it}\end{aligned}$$

where α is a fixed intercept, α_i is a random intercept of state to control initial heterogeneity among states, $\boldsymbol{\beta}$ is a 1×11 vector to represent the coefficients of confounding variables, and the smoothing function f(t) is a penalized spline for calendar time t (1 = 1999, …, 19 = 2017) to take temporal autocorrelation into account. Because overdispersion existed in the model (i.e., $E(Y_{it}) < \mathrm{Var}(Y_{it})$), we used quasi-likelihood estimation to solve the problem. All smoothing parameters in our model were estimated with the generalized cross-validation method to provide a greater accuracy, especially in cross-basis functions. The effect of gun law provisions along with lags was mainly evaluated in the cross-basis function $f(\mathrm{LAW}_{it}, L = k)$, which is an interaction term between a basis function for the number of gun law provisions (LAW_{it}) and another basis function for lag (L). The basis function was defined as a natural cubic spline. Notice that the cross-basis function must generate missing data in the first k observations in the gun law variable in each state when the maximum lag is equal to k; thus, we selected the maximum lag by 8 in this study because the generated missing gun laws in the first 8 years (i.e., 1999–2007) can be imputed by the gun law data from 1991 to 1998. This operation can efficiently prevent the reduction of statistical power caused by missing data. Lastly, the term offset_{it} is the logarithm of population in state i and calendar time t.

Table 1
The original regulation categories defined by the State Firearm Laws website and the merged regulation groups in this study.

Original regulation categories[b]	Merged regulation groups
Buyer regulations ($n = 17$)	Buyer regulation group ($n = 28$)[a]
Background checks ($n = 11$)	Dealer regulation group ($n = 28$)[a]
Dealer regulations ($n = 17$)	
Domestic violence ($n = 21$)	Domestic violence regulation group ($n = 21$)
Assault weapons and large-capacity ammunition magazines ($n = 8$)	Gun types, components, and trafficking regulation group ($n = 22$)
Ammunition regulations ($n = 7$)	
Gun trafficking ($n = 7$)	
Prohibitions for high-risk gun possession ($n = 11$)	Possession, access, and carrying regulation group ($n = 41$)
Possession regulations ($n = 12$)	
Child access prevention ($n = 11$)	
Concealed carry permitting ($n = 7$)	

n = The number of laws in a category.
[a] The background check regulations were merged with both buyer regulations and dealer regulations because the background check provisions primarily centered on commercial transactions related to firearm sales.
[b] The regulations of immunity, preemption, and stand your ground were excluded because they did not directly regulate access and availability of firearms and firearm products in contrast to the other provisions that we focused on this study.

All estimated coefficients for linear terms (i.e., $\widehat{\boldsymbol{\beta}}$) can be transformed into RR of firearm homicide by an exponential function. The estimated cross-basis function can be also transformed into RR along with the number of gun law provisions and lag, where the reference levels of gun law provision amount and lag are 0s. In particular, we can specify the value of lag (i.e., L) in the estimated cross-basis function to explore how many years a state with a specific number of enacted gun law provisions may significantly reduce firearm homicide mortality. Lastly, by accumulating the estimated cross-basis function across all lags, we can evaluate the overall influence between gun law provisions and firearm homicide mortality. The concern of multicollinearity among confounding variables was also examined and disposed of without further actions because tolerances were all larger than 0.1. Both time smoothing function and random effect were significant in our models with different gun law measures.

The data cleaning and management were accomplished by SAS, version 9.4 (SAS Institute, Cary, NC), and all data analyses were performed in R software package, version 3.4.0 (R Development Core Team, Vienna, Austria). The significance level was determined by 5%.

Results

Table 2 shows that, from 1999 to 2017, on average, a state had 24.91 (standard deviation [SD] = 23.85) state gun law provisions in place per year. The largest number of total gun law provisions (106) appeared in California in 2017. Among the five regulation groups, the possession, access, and carrying group had 11.01 (SD = 6.88) gun law provisions on average, which is much larger than the averages of the other four groups. Moreover, every year a state had 237.44 (SD = 291.86) firearm homicide deaths, on average. After taking state population into account, the average crude firearm homicide rate for a state was 3.39 (SD = 2.13) persons per 100,000 people.

The contour plot in Fig. 1 presents the variation of RR of firearm homicide mortality along with gun law provisions and lags, where the blue cold spot represents an RR less than 1. No matter the total provisions or regulation groups, more gun law provisions result in smaller RRs of firearm homicide after lag 5 or lag 6. Considering all gun law provisions in each lag, Fig. 2 reveals that when the total number of gun law provisions in a state increased, a significant RR less than 1 was not observed until lag 7. The minimum RR was 0.52 (95% confidence interval [CI] = 0.35, 0.79) at lag 8 when a state has 106 gun laws in place.

Considering regulation groups by lags, most of them did not result in significant findings in shorter lags from 1 to 4, except for the domestic violence regulation group (lag = 3 and 4, Fig. 3c), gun types, components, and trafficking group (lag = 4, Fig. 3d), and possession, access, and carrying regulation group (lag = 1, Fig. 3e); however, significantly reduced RRs more likely appeared in longer lags from 5 to 8. Fig. 4a indicates that a significant RR less than 1 was observed at lag 7 and lag 8 when the buyer regulation group had more than 20 related provisions. Significant reductions in the RR were obtained since lag 6 in the dealer regulation group, refer

Table 2
Summary statistics of gun law provisions by group and firearm homicide in the United States, 1999–2017.

Variable	Mean	SD	Min	Q1	Median	Q3	Max
Total gun law provisions	24.91	23.85	2	10	15	29	106
Buyer regulation	4.29	6.16	0	0	1	7	23
Dealer regulation	4.69	6.75	0	0	1	8	25
Domestic violence regulation	3.44	4.82	0	0	1	5	18
Gun types, components, and trafficking	1.84	3.60	0	0	0	2	18
Possession, access, and carrying	11.01	6.88	0	7	9	14	33
Firearm homicide deaths	237.44	291.86	0	30	124.50	347	1883
Crude firearm homicide rate[a]	3.39	2.13	0	1.60	3.19	4.80	11.40

SD = standard deviation; Q1 = 1st quartile; Q3 = 3rd quartile.
[a] The unit of rate is per 100,000 populations.

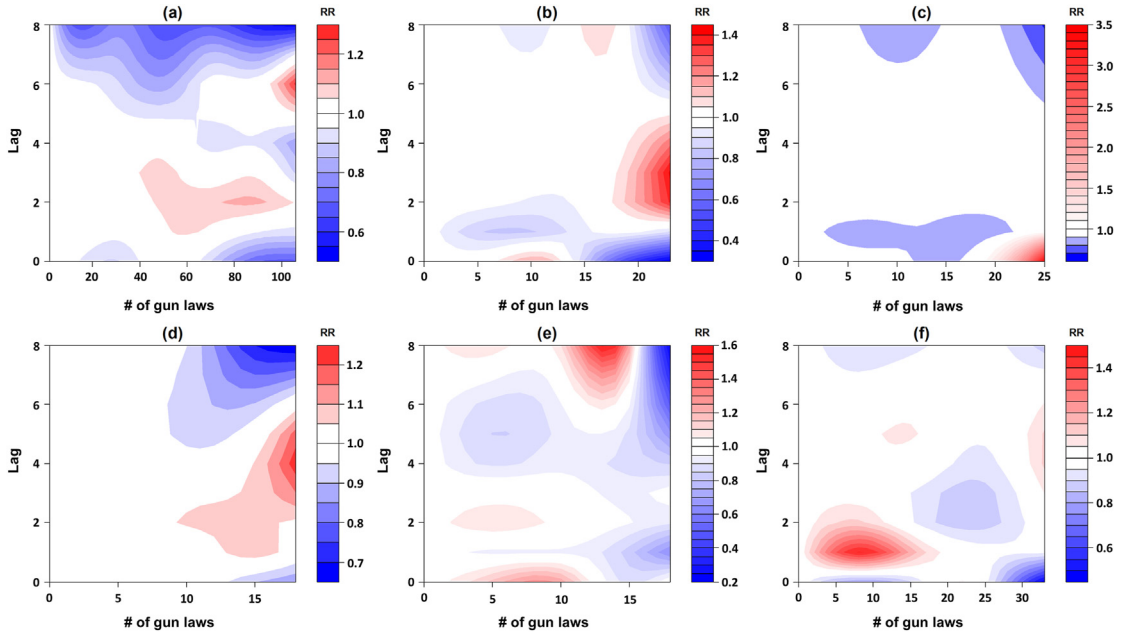

Fig. 1. The variation of relative risk of firearm homicide mortality by lags and (a) total gun law provisions, (b) buyer regulation group, (c) dealer regulation group, (d) domestic violence regulation group, (e) gun types, components, and trafficking regulation group, and (f) possession, access, and carrying regulation group.

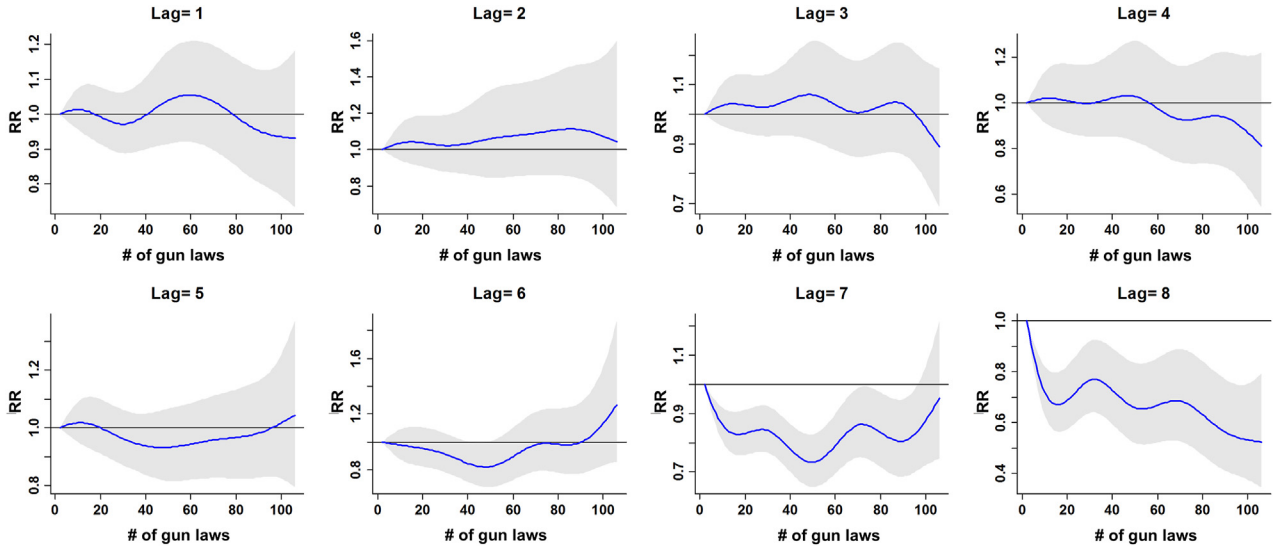

Fig. 2. The variation of relative risks of firearm homicide mortality and 95% confidence intervals (gray areas) along with the increase of total gun law provisions by lag.

Fig. 4b. Fig. 4c reveals that when a state had at least 10 gun law provisions in the domestic violence group, a significant RR less than 1 was observed from lag 5 to lag 8. The gun types, components, and trafficking group also had a negative association with firearm homicide mortality with a longer lag, especially with at least 15 related gun law provisions, refer Fig. 4d. Among the five groups, the lowest significant RR = 0.24 (95% CI = 0.15, 0.39) was observed in this group at lag 8 when the number of related gun law provisions reached 18. Significant findings were only found at lag 7 in the possession, access, and carrying group (Fig. 4e).

For the overall influence of gun law provisions on firearm homicide, Fig. 5a shows that, when a state had a total of 97 gun law provisions in place, the RR reduced to the lowest level by 0.30 (95% CI = 0.22, 0.42). In the buyer regulation group, the RR was significantly declined as low as 0.21 (95% CI = 0.13, 0.35) with 23 gun law provisions in place, refer Fig. 5b. However, more gun laws in the dealer regulation group did not necessarily demonstrate a reduced RR, where the overall effect of the dealer regulation group had a significant RR less than 1 only up to 18 gun law provisions (RR = 0.58; 95% CI = 0.74, 0.95). The domestic violence group needed to accumulate at least 6 gun law provisions to reduce RR significantly less than 1 (RR = 0.92; 95% CI = 0.85, 0.99), refer Fig. 5d. In the gun types, components, and trafficking regulation group, Fig. 5e shows that the significant negative association between gun laws and firearm homicide mortality did not appear until 15 related gun law provisions were in place (RR = 0.64; 95% CI = 0.49, 0.84). Lastly, a state should have at least 21 gun law provisions related to the possession, access, and carrying group to

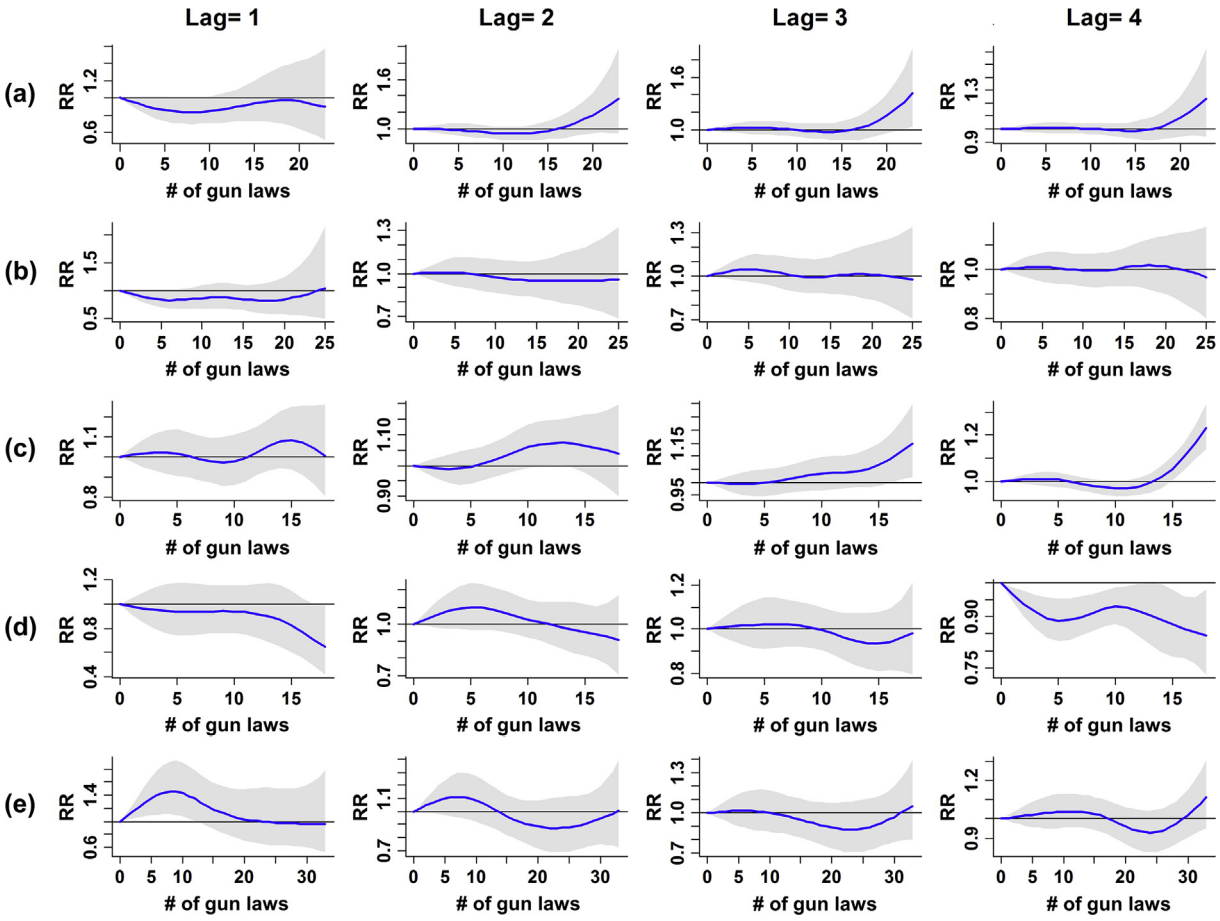

Fig. 3. The variation of relative risks of firearm homicide mortality and 95% confidence intervals (gray areas) along with the increase of gun law provisions from lag 1 to lag 4 in the (a) buyer regulation group, (b) dealer regulation group, (c) domestic violence regulation group, (d) gun types, components, and trafficking regulation group, and (e) possession, access, and carrying regulation group.

significantly reduce the RR of firearm homicide mortality less than 1, refer Fig. 5f.

Among the examined confounding variables, the law enforcement officer rate is the only factor that appeared to be significantly and negatively associated with firearm homicide mortality in all groups. For more details, refer to Table 1S in Appendix.

Discussion

Gun violence is a public health problem in the U.S. Federal law establishes nationally applicable gun regulations, but these regulations do not cover all instances of gun acquisition or possession.[6] States have also taken approaches to regulate gun acquisition and possession, resulting in state-level variation in gun laws and their implementation.[7] This study uses a novel statistical modeling approach to explore relationships between state gun law provisions and health data. It confirms that the reduction of firearm homicide mortality may not be anticipated in the presence of generic state gun laws in the short term. On average, states had to wait at least 7 years to see a significant risk reduction. In terms of specific gun law regulation groups, the significant lagged period can be shortened to 4 or 5 years. Moreover, having more gun law provisions in the gun types, components, and trafficking group seems to reduce the RR of firearm homicide mortality to the lowest level, compared with the other regulation groups. Across groups, more related gun law provisions eventually appear to significantly reduce firearm homicide mortality, with the exception of the domestic violence group. This study provides additional evidence to confirm that state gun laws may contribute to reducing gun violence and that implementation over time is a critical variable.

Rather than analyzing individual gun laws similar to some previous studies,[25] this study used an immediate accumulation of the existing gun laws. The rationales to consider gun laws in an accumulated fashion are as follows: First, most gun laws are highly correlated and are inadequate to be analyzed independently in the same model; second, legal provisions may be part of a 'package' of gun law reforms; third, fitting highly correlated gun laws in the same model could bias results because of multicollinearity.[26] Adopting this strategy to quantify the strength of gun laws has pros and cons. A recent trend analysis relied on these same gun law provision measurements to reveal state gun law disparities.[27] However, direct summation assumes a fair impact of each gun law on firearm homicide, although the literature has shown diverse health impacts among different gun laws.[7] Various scoring systems have been created and used to assess the adequacy of gun law provisions, while those scores were not updated every year.[28] Applying variable condensation might be a solution, but the association between gun violence and a condensed variable does not have a truly meaningful explanation because the condensed variable has lost a quantified unit. Thus, while imperfect, the number of gun law provisions is a useful proxy because it combines all gun laws in place into one variable, which can both prevent the multicollinearity problem and stand in for the magnitude of gun law reform efforts. This study considered the potential impact of laws

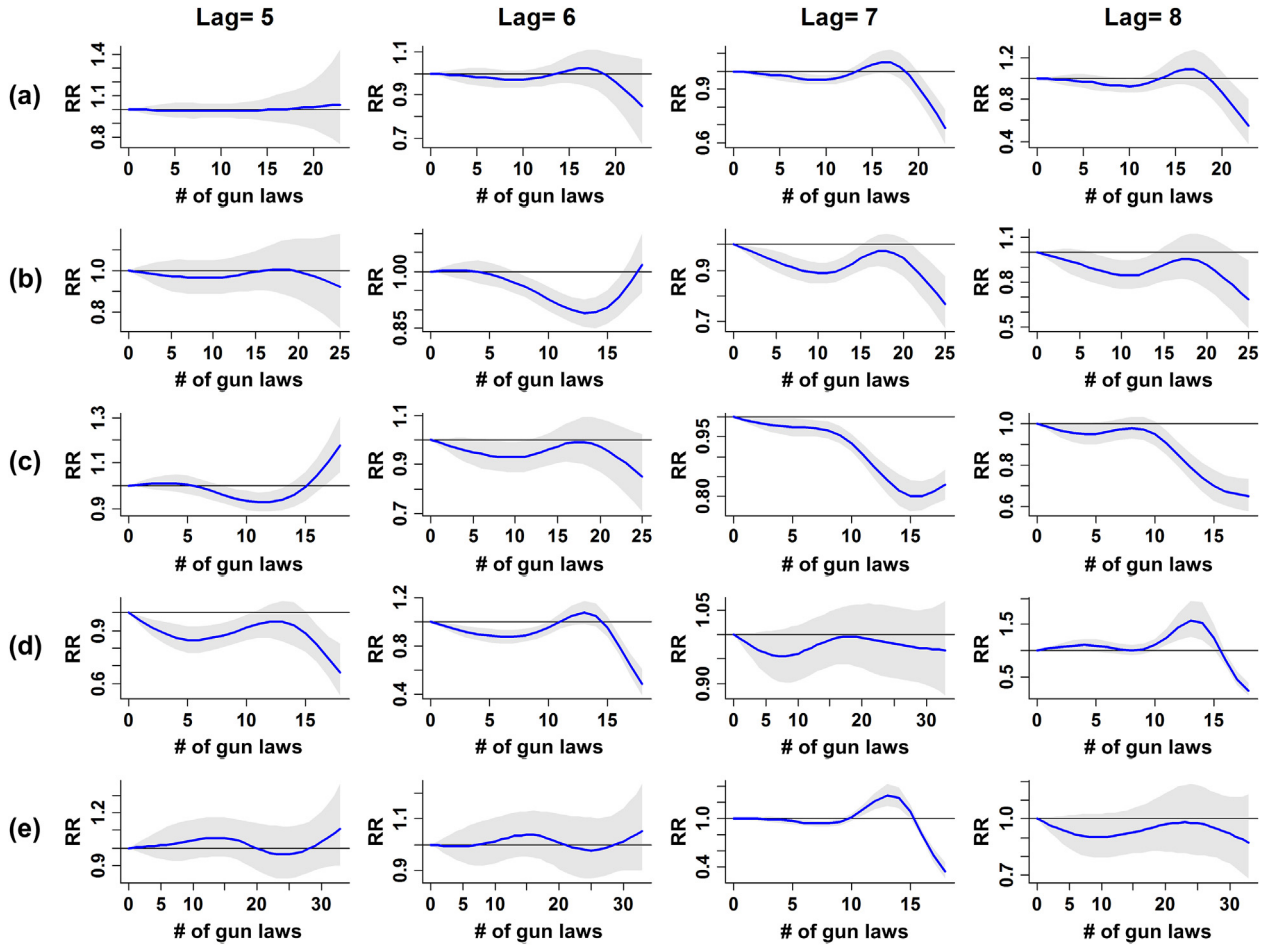

Fig. 4. The variation of relative risks of firearm homicide mortality and 95% confidence intervals (gray areas) along with the increase of gun law provisions from lag 5 to lag 8 in the (a) buyer regulation group, (b) dealer regulation group, (c) domestic violence regulation group, (d) gun types, components, and trafficking regulation group, and (e) possession, access, and carrying regulation group.

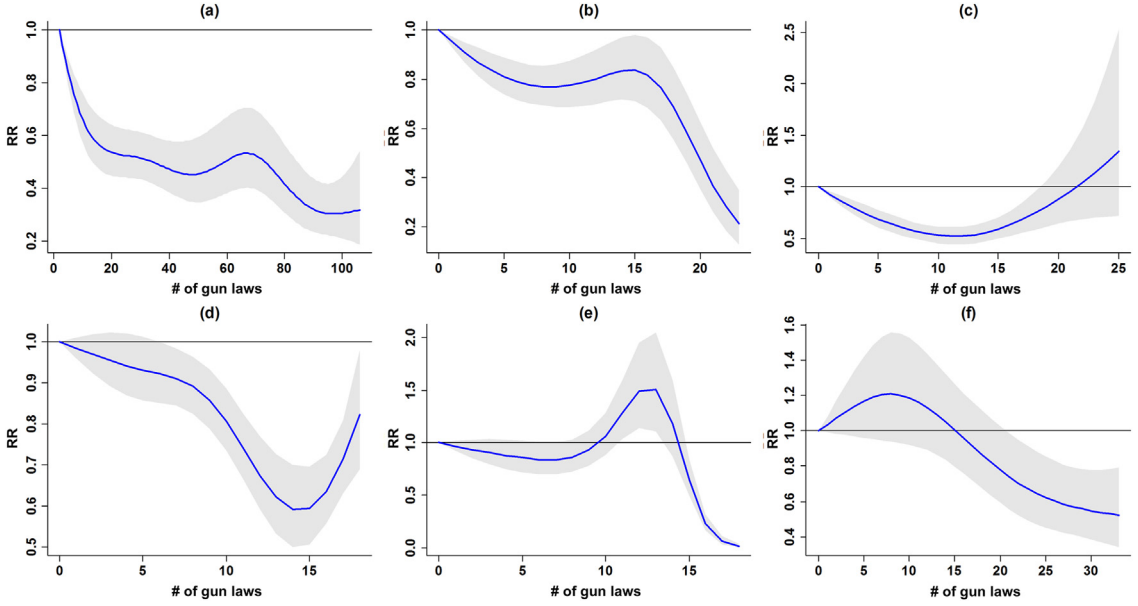

Fig. 5. Overall influence of state gun law provisions on firearm homicide mortality in the (a) total gun laws, (b) buyer regulation group, (c) dealer regulation group, (d) domestic violence regulation group, (e) gun types, components, and trafficking regulation group, and (f) possession, access, and carrying regulation group.

over periods of time because the connection between adopted laws and health outcomes involves implementation practices and other changes that do not occur immediately but operate over stretches of time.[17] This analysis aimed to pinpoint periods of time of meaningful change to homicide rates connected to these laws so that such periods could be examined further to identify specific factors contributing to these changes.

Our findings that the RRs of homicide were different by groups of provisions suggest that certain provisions should be examined further to determine whether they may be more impactful than others. For instance, provisions that focus on gun buyers and domestic violence may be especially promising strategies to address homicide. Groups that showed consistent and meaningful reductions in the RR of homicide as the number of provisions increased and over time, such as the buyer regulation and domestic violence regulation groups, merit particular attention. These groups may contain specific provisions that, as enforced, most meaningfully impact homicide. Pinpointing such provisions and examining their enforcement could help identify which laws on the books, as implemented, are most critical from a homicide prevention standpoint. Similarly, the lowest reduction to the RR that we observed in the possession, access, and carrying group suggests that provisions in this group may face implementation hurdles or may need revisiting if homicide prevention is their ultimate goal.

Studying and understanding enforcement is difficult because there may be insufficient data to determine whether, when, and how a gun law was enforced or adequate statistical methods. This study applied a novel model to examine if the effectiveness of state gun law provisions changed over time. Its finding that effectiveness could diminish over time emphasized the importance of sustained gun law efforts and implementation.

The strength of this study is in using longitudinal gun law and gun violence data. The most apparent limitation is focusing purely on the number of state laws on the books, which does not consider enforcement or account for local variation. Therefore, we initially assumed that each gun law provision had an equal impact on firearm homicide. The in-state variation of firearm homicide death or potential confounding variables at the county level is also not considered. This may explain, for instance, why we found the law enforcement officer rate to be a confounding variable that is significantly and negatively associated with firearm homicide mortality. Historically, findings in the literature are mixed about the relationship between numbers of officers and crime rates, with recent analyses concluding that any relationship that may exist is at least small and possibly not terribly meaningful.[29,30] Mixed findings are echoed in the homicide context, where no consensus has existed about whether the number of officers meaningfully impacts homicide rates, but recent findings suggest an expanded police force is connected to declining rates of homicide.[31] Our finding is consistent with this recent work. Alternatively, it could be the product of other differences within states or our reliance on state-level data, which may obfuscate nuanced connections between localized, per capita officer and homicide rates. Furthermore, our data did not include non-fatal gun injuries and suicides, so the overall impact of gun laws might be underestimated. Finally, the distribution lag non-linear model does not propose a way to examine the goodness of fit, so the adequacy of the model fitting is uncertain.

Conclusions

State regulation of gun acquisition and possession varies and operates with the backdrop of uniform federal laws that apply only to some instances of gun acquisition and possession. Given this state-by-state variability, understanding what happens to homicide rates over time and as state laws change is critical. This study demonstrated that even in states with a certain number of gun laws, there appears to be a lag period of at least 7 years, on average, before the risk of firearm homicide declines. In addition, groups of provisions that regulate various aspects of firearms appear to have different relationships to firearm homicide rates. Examining the impacts of different types of state gun policy reforms and their implementation over time can help understand the effectiveness of gun laws and should be the focus of future research.

Author statements

Ethical approval

Not applicable.

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Competing interests

None declared.

Patient consent

Not applicable.

Data sharing

The raw data are available from the Centers for Disease Control and Prevention Wide-ranging Online Data for Epidemiologic Research and State Firearm Laws websites.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2020.08.028.

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