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Original Research

A lesson from history? Worsening mortality and the rise of the Nazi Party in 1930s Germany

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ABSTRACT

Objectives: The aim of the study was to test the hypothesis that worsening mortality rates in the early 1930s were associated with increasing votes for the Nazi Party.
Study design: The study consist of panel data with fixed effects.
Methods: We used district- and city-level regression models of Nazi vote shares on changes in all-cause mortality rates in 866 districts and 214 cities during federal elections from 1930 to 1933, adjusting for election and district/city-level fixed effects and sociodemographic factors. As a falsification test, we used a subset of deaths less susceptible to sociopolitical factors.
Results: Historical downward trends in mortality rates reversed in the early 1930s in Germany. At the district/city level, these increases were positively associated with a rising Nazi vote share. Each increase of 10 deaths per 1000 population was associated with a 6.51-percentage-point increase in Nazi vote share (95% confidence interval = 1.17–11.8). The strongest associations were with deaths due to infectious and communicable diseases, suicides, and alcohol-related deaths. Worsening mortality had no association with votes for the Communist Party or for other contemporary political parties. Greater welfare payments were associated with smaller increases in both mortality and Nazi vote share, and adjusting for welfare generosity mitigated the association by approximately one-third.
Conclusions: Worsening mortality rates were positively associated with the rise of the Nazi Party in 1930s Germany. Social security mitigated the association between mortality and Nazi vote share. Our findings add to the growing evidence that population health declines can be a ‘canary in the coal mine’ for the health of democracies.

Introduction

Although the Weimar Republic (1918–1933) was a short parenthesis in Germany's history, it is among its most consequential. The rise of Nazism in the early 1930s, in the wake of the Great Depression, coincided with worsening economic hardship[1] characterized by severe austerity policies,[2] mass unemployment,[3] and widespread discontent. Yet, while other countries also faced economic insecurity, Germany experienced an increase in mortality starting in 1931–1932, at a time when mortality was declining in other European nations including Poland, France, and the Netherlands.

A series of recent articles have found a striking and consistent correlation between deteriorating population health and support for populist radical right parties.[4] Bor,[5] Wasfy et al.,[6] Herrin et al.,[7] and Bilal et al.[8] found that those counties in which life expectancy stagnated or declined from 1980 to 2014 exhibited substantially higher vote shares for Donald Trump in the 2016 presidential election. Similar patterns were observed in the UK, where worsening mortality, alongside budget reductions, was positively associated with greater votes for Brexit.[9,10] Several studies have linked ‘deaths of despair,’ including external causes of death such as suicides and alcohol-related mortality, to a greater share of votes for the Republican Party.[11–13] Stagnating or rising mortality has been linked to economic dislocation in the absence of a robust safety net. US areas that lost manufacturing jobs to Chinese competition in the

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2000s saw increasing male midlife mortality and a shift among voters toward Republican candidates in US House and Presidential elections.[14,15]

Support for the Nazi Party rose from 2.68% of the German electorate in 1928 to 43.9% by 1933. However, the rise in Nazi support varied geographically. Here, we test the hypothesis that worsening mortality rates at the state, district, and city level were associated with the rise of the Nazi Party in Germany, even after accounting for economic hardship and other potential confounding factors. Specifically, we test whether the increase in mortality in the early 1930s was a 'canary in the coal mine' for the growth in Nazi support, asking whether populations suffering health declines are fertile ground for populist politics.

Methods

Sources

Voting data were obtained for four federal elections between 1930 and 1933 (September 1930, July and November 1932, and March 1933) covering 16 states, 866 districts, and 214 cities. Details were described elsewhere,[2] and we provided full details of our sources in a supplementary file, but briefly, these were drawn from official German statistics. In the German federal elections, under the principle of proportional representation, people voted for nationwide party lists to decide who would be the chancellor. To measure support for the Nazi Party, we captured vote shares as a proportion of the total vote. We also measured changes in other parties' shares, including the Social Democratic Party, the Center Party, the Communist Party, and a residual category of other small parties. We evaluated changes in vote shares across elections (e.g., between 1930 and 1932 and from 1932 to 1933) and over the entire period (1930–1933).

To measure mortality rates, we used three levels of aggregation based on the data characteristics. We extracted new data on the main vital statistics (population, births, and deaths) at the district level, covering the calendar years 1927–1933 ($n = 866$). We also collected mortality data for 214 cities, where cause-specific mortality was available for 20 causes of death (city-level data being available for 1928, 1932, and 1933). In addition, at the state level, we disaggregated annual mortality data into 7 age bands to calculate age-standardized mortality rates for each German state ($n = 16$). Annual crude death rates are calculated per thousand population. At the city level, we further disaggregated deaths into two groups. One includes deaths plausibly linked to deteriorating social conditions over the short term (including deaths from suicides, homicides and murders, and stroke), whereas we used a second group as a falsification test, which should be less responsive to sociopolitical changes including cancer deaths.

To adjust for economic decline, based on district-level tax returns, we estimated per capita (taxable) income in each district. At the city level, we proxy economic conditions by the number of unemployed as a share of the total population. Finally, we also proxy the extent to which the Weimar Republic provided social protection, using newly extracted district-level data from official sources on welfare spending per capita, capturing social security payments from central governments and lower levels of government on open care (mostly relief and medical assistance). Appendix A shows descriptive statistics for all variables.

Methods

Multivariate regression models were used to quantify the association between German mortality rates and Nazi vote share using a difference-in-differences with an intensity of treatment interpretation based on:

$$\mathrm{NAZI}_{d,t} = \alpha + \beta_1\,\mathrm{Mortality}_{d,t\text{-}1} + \beta_2\,\mathrm{Income}_{d,t\text{-}1} + \beta_3\,\mathrm{Welfare}_{d,t\text{-}1} + \gamma_d + \delta_t + e_{d,t} \qquad (1)$$

where d denotes districts, t is one of the four election years (September 1930, July and November 1932, and March 1933), and $\mathrm{NAZI}_{d,t}$ is the vote share of the Nazi Party in percentages of the total vote (%). Because age-standardized mortality data are not available at the local level, $\mathrm{Mortality}_{d,t\text{-}1}$ is simply the crude death rate in the year before the election year, $\mathrm{Income}_{d,t\text{-}1}$ denotes income per capita, and $\mathrm{Welfare}_{d,t\text{-}1}$ denotes the welfare payments per capita in the year before the election, both expressed in nominal terms. All models include district-level fixed effects (γ_d) and fixed effects for the calendar years of 1932 and 1933 (δ_t). By pooling data for different elections and using time and district fixed effects, mortality here can be interpreted as excess mortality or deviations of mortality from the within-district sample mean. Finally, $e_{d,t}$ is the error term. Standard errors are clustered at the state level, although clustering at lower levels (i.e., district) displays the same levels of statistical significance. For simplicity, when we say states, we also mean Prussian provinces.

Results

Association of mortality with Nazi voting

We first investigate the visual association of mortality increases and Nazi voting shares at the state level, followed by subsequent analyses at the district and city levels.

Fig. 1 depicts the unconditional association between the change in state mortality rates from 1928 to 1933 and the Nazi vote share in 1933 at the state level. As shown, those states with greater mortality rises also experienced larger votes for the Nazi Party in 1933 ($r = 0.41$, P-value $= 0.00$). A similarly strong correlation was evident at the district level ($r = 0.11$, P-value $= 0.00$).

Table 1 shows the results from multivariate regression models on the impact of mortality and the Nazi Party vote share estimated at the district level and sequentially adjusted for economic hardship (measured by income per capita). Fixed effects would have adjusted for any time-invariant characteristics of the districts (possibly, environmental and sociodemographic confounders). Even after adding these controls, worsening health was associated with rise in Nazi votes. When we use data for the four elections

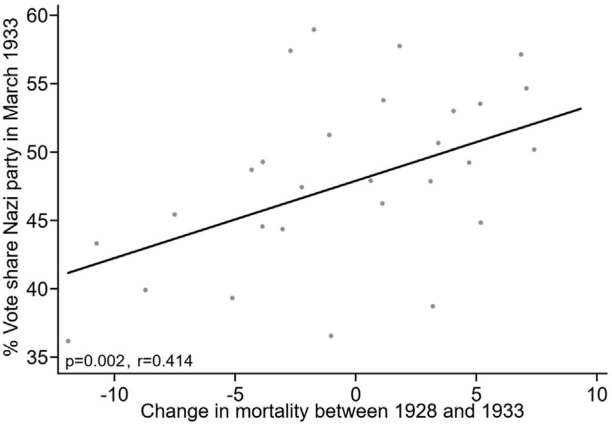

Fig. 1. Percentage of state-level vote shares for the Nazi Party in 1933 and changes in the crude death rate between 1928 and 1933.

Table 1

District-level impact of mortality on the Nazi Party vote share.

Variable	Elections, 1930 and 1933 (n = 1732)		Elections, 1930 and 1932 (both) (n = 2598)		All elections (n = 3464)	
	b (95% CI)		b (95% CI)		b (95% CI)	
Mortality	0.88	(0.23, 1.53)	0.78	(0.11, 1.45)	0.65	(0.12, 1.18)
Income	−2.88	(-3.88, −1.87)	−2.34	(-3.35, −1.33)	−2.32	(-3.36, −1.27)

CI = confidence interval.

Parameter estimates are unstandardized regression coefficients representing the percentage point increase in vote shares for the Nazi Party associated with an increase of 1 death per 1000 population. We use a balanced panel of 866 districts and pooled data for different elections. All models control for income per capita and add district-level fixed effects and a fixed effect for to the election time: a 1929–1930 fixed effect for the 1930 election, a 1931–1932 fixed effect for the elections of 1932 (either June or November), and a 1932–1933 fixed effect for the election of 1933. Robust standard errors are clustered at the state level (Prussian provinces are considered different states).

between 1930 and 1933, each increase of 10 deaths per 1000 population was associated with a 6.51-percentage-point increase in the Nazi vote shares (95% confidence interval [CI] = 1.17 to 11.8). The results were robust to focusing on just the 1930–1932 elections and the 1932–1933 elections.

Next, at the city level, we disaggregated all-cause mortality into those types that could plausibly be impacted by changes in contemporary social conditions, including suicides, homicides, and stroke, whereas a second group, which should be less responsive in the short term to sociopolitical changes including cancer deaths, was used as a falsification test. Table 2 shows the results of these city models. For those linked to social conditions, each increase of 10 deaths per 1000 population was associated with a 31.71-percentage-point increase in the Nazi vote shares (95% CI = 11.91 to 51.50). By way of contrast, when we compare these patterns with causes of death that are less responsive to short-term social conditions, using cancer mortality, the results are not statistically significant (13.10; 95% CI = −12.42 to 38.61). Appendix B reports disaggregated models for each of the 20 causes of death, revealing that Nazi vote rises were most closely associated with infectious and communicable diseases such as tuberculosis, measles, and lung infections, as well as the aforementioned deaths from suicides, alcohol-related deaths, and stroke.

Comparing vote patterns by district characteristics

In Appendix C, we explore potential heterogeneity in the association of mortality increases and the Nazi electoral boost, stratifying by district characteristics. Using data from the census of 1933, we split the sample for districts with occupational and religious group shares below and above the median value. When we stratify the sample, we find that mortality had a stronger association with Nazi vote shares in more agricultural districts and a weaker association in more industrial districts. This finding is consistent with the prior literature that found that the unemployed did not favor the Nazis at the polls (they turned to the Communists).[2,3] Instead, it was those just above in the economic hierarchy, who also experienced great hardship, who had something to lose, who favored the Nazis. The effects of mortality were also stronger in areas with a higher number of Protestants and Jews. With respect to the latter, Voigtländer and Voth[16] outlined the importance of long-standing antisemitism in the Nazi electoral boost.

Table 2

City-level impact of amenable and non-amenable mortality on the Nazi Party vote share.

Variable	Elections, 1928 and 1933 (n = 428)	
	b (95% CI)	
Amenable mortality	3.17	(1.19, 5.15)
Non-amenable mortality	1.31	(-1.24, 3.86)

CI = confidence interval.

Parameter estimates are unstandardized regression coefficients representing the percentage point increase in vote shares for the Nazi Party associated with an increase of 1 death per 1000 population. We use a balanced panel of 214 cities and pooled data for two elections. Amenable mortality groups the following causes: stroke, suicide and murder, or homicide. Non-amenable mortality includes deaths from cancer. We add city-level fixed effects and a 1932–1933 fixed effect for the election of 1933. Although cause-specific mortality is only available in 1928 and 1933, the comparison from 1928 to 1933, coming from local deviations from undisturbed levels, is a good measure from the German suffering in the elections of March 1933. Robust standard errors are clustered at the state level (Prussian provinces are considered different states).

Mitigating role of welfare payments

We further evaluated the role of welfare payments to buffer economic suffering, which may have acted as a confounder. Appendix D shows that adding a control for welfare payments significantly reduces the impact of mortality on the Nazi vote share. The results are still statistically significant for mortality in some models. However, the coefficient on welfare payments is also statistically significant, and its negative sign indicates that social security may have mitigated radicalization. Importantly, for our argument, the size of the coefficient for mortality when controlling for welfare is cut down by between one-third and two-thirds based on specification, for instance, by 4.85 deaths per 100,000 population (95% CI: 1.37 to 8.33), when considering data for the four elections between 1930 and 1933.

Robustness checks

We performed a series of robustness tests. As shown in Table 1, we used data from different elections, showing that our results are not specific to a single election, but to a process of worsening social and financial conditions that collapsed in the early 1930s. In Appendix E, instead of using panel data with fixed effects, we show consistent results computing the percentage point change in the crude death rate (the change in the level) between elections. In Appendix F, we also show consistent results using infant mortality rates instead of crude death rates for all ages, despite having a lower predictive value, probably because births to the poorest families fall disproportionately during a recession.[17] Although disaggregated age bands were not available at the district or city level, these were available at the state level for 7 age bands. Despite state-level models being statistically underpowered (relying on 16 data points for 16 states as in the age-band statistics Prussian provinces counted as a single state), we found that by using the age bands to calculate age-standardized mortality rates, none of the results was qualitatively changed (see Appendix G). Using Eq. (1), a further check we did in the age structure was interacting the mortality rates with time dummies (Appendix H). These interaction terms ensured that our models are not simply showing worsening health of older Germans at a time of crisis. Finally, to further account for falling revenues as a result of the Great Depression, we also show consistent results with those in Table 1 after interacting a precrisis income indicator (in 1928) with time dummies (Appendix I). Our bottom line is that the association between mortality and the rise in Nazi support persists even after adjusting for income and other sociodemographic factors, implying that poor health is not simply a proxy or mediator for a relationship between income and Nazi support.

Discussion

Our analysis shows a significant association between mortality rates and increasing vote shares for the Nazi Party in 1930s Germany. These rises in Nazi vote share were most strongly correlated with deaths from infectious diseases including tuberculosis and pneumonias, as well as suicides, homicides, and heart attacks. In addition, we found that adjusting for levels of welfare support attenuated this correlation, suggesting that the generosity (or stinginess) of state and local welfare regimes may have been a common cause of the observed health and political trends.[18,19]

Study limitations

Before interpreting our findings further, we must note several important limitations. First, we were unable to adjust for age at district and city levels, creating potential for error. However, our results were consistent when performed at the state level, wherein direct age adjustment was possible. Furthermore, our interaction terms and district- and city-level fixed effects would have adjusted for any time-invariant characteristics of the age distribution. Second, it is possible that a third underlying factor drove both mortality rises and Nazi vote shares. Indeed, we found that adjusting for welfare generosity attenuates the mortality-voting relationship. Hence, the observational analysis can only demonstrate correlation, rather than causal effects. However, this modeling approach with controls and fixed effects follows that of recent analyses on mortality and Trump votes in the US[5] and on mortality and Brexit votes in the UK.[9] Whether or not they are a direct cause, mortality increases appear to be an early-warning measure for political polarization.

Our article has implications for future research. Welfare payments seem important in attenuating the Nazi-mortality association, which is consistent with the historical literature that budget cuts implemented in 1931–1932 by Chancellor Brüning, commonly known as the 'Hunger Chancellor,' positively correlated with the rise of the Nazi Party.[2] We also showed that the association between mortality rates and Nazi vote shares was strongest in rural districts. Yet, further research is needed to investigate these underlying mechanisms and better understand the complex causal chains involved.

In Appendix J, we show that the Nazi Party was the only party that managed to transform German suffering into more votes. Neither the communists, who traditionally were seen as guardians of the interests of working people, the Social Democrats, the political home of the workers' movement and middle classes, nor the Center party, a conservative catholic party, saw gains in support with the declining health of the electorate. We interpret this as evidence that at times when people are suffering, they may be more open to the siren calls of right-wing radical populist parties.

Our article supports the notion that epidemiological data can serve as a 'canary in the coal mine,' identifying populations that are being left behind by social progress, which may in turn create fertile ground for receptivity to populist messages. As Keynes gloomily foreshadowed in *The Economics Consequences of the Peace*, "If we take the view that for at least a generation to come Germany cannot be trusted with even a modicum of prosperity … that year by year Germany must be kept impoverished and her children starved and crippled … If we aim deliberately at the impoverishment of Central Europe, vengeance, I dare predict, will not limp."[20](p199-200).

Author statements

Ethical approval

The authors have no conflicts to declare. The authors used historical statistical records based on official data.

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Competing interests

The authors have no conflicts to declare.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2021.03.022.

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Review Paper

Cervical cancer screening uptake in Sub-Saharan Africa: a systematic review and meta-analysis

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ABSTRACT

Objectives: The objective of this study is to estimate the pooled uptake of cervical cancer screening and identify its predictors in Sub-Saharan Africa.
Study design: Systematic review and meta-analysis.
Methods: We searched PubMed, EMBASE, CINAHL, African Journals OnLine, Web of Science and Scopus electronic databases from January 2000 to 2019. All observational studies published in the English language that reported cervical cancer uptake and/or predictors in Sub-Saharan Africa were initially screened. We assessed methodological quality using the Newcastle-Ottawa Scale. An inverse variance-weighted random-effects model meta-analysis was performed to estimate the pooled uptake and odds ratio (OR) of predictors with a 95% confidence interval (CI). The I^2 test statistic was used to check between-study heterogeneity, and the Egger's regression statistical test was used to check publication bias.
Results: We initially screened 3537 citations and subsequently 29 studies were selected for this review, which included a total of 36,374 women. The uptake of cervical cancer screening in Sub-Saharan Africa was 12.87% (95% CI: 10.20, 15.54; $I^2 = 98.5\%$). A meta-analysis of seven studies showed that knowledge about cervical cancer increased screening uptake by nearly five times (OR: 4.81; 95% CI: 3.06, 7.54). Other predictors of cervical screening uptake include educational level, age, Human Immune deficiency Virus (HIV) status, contraceptive use, perceived susceptibility and awareness about screening locations.
Conclusions: Cervical screening uptake is low in Sub-Saharan Africa as a result of several factors. Health outreach and promotion programmes to target these identified predictors are required.

Introduction

Cervical cancer is a global public health challenge.[1] The primary cause of cervical precancer and cancer is persistent infection with one or more of the high-risk oncogenic types of human papillomavirus (HPV). HPV interferes with the normal functioning of cells, which results in distinct changes in the epithelial cells of the transformation zone of the cervix.[2] Cervical cancer is one of the very few types of cancers where a precancer stage lasts for many years before becoming invasive cancer, thus allowing ample opportunity for detection and treatment.[3] Cervical cancer is a malignancy for which effective screening is available. The screening seeks to identify precancerous cellular changes on the cervix that may become cervical cancer if they are not appropriately treated.[4]

Cervical cancer is the fourth most common cancer in women, with an estimated 530,000 new cases every year, representing 7.9% of all female cancers.[5] In 2015, approximately 90% of the 270,000

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deaths from cervical cancer occurred in low- and middle-income countries.[5] The mortality rate varies remarkably among different regions of the world, with rates ranging from <2 per 100,000 in Western Europe and New Zealand to 27.6 per 100,000 in Sub-Saharan Africa.[6]

Cervical cancer prevention and the impact of screening programmes on cervical cancer-related deaths have been given considerable attention in developed countries, in contrast to the minimal effort seen in most low- and middle-income nations.[7] Cervical cancer screening coverage is very limited in low- and middle-income countries, as shown by a study that reported coverage of cervical cancer screening in developing countries to be 19% (on average) compared with 63% (on average) in developed countries.[8] Data from the 2017 World Health Survey indicated that the coverage of cervical cancer screening was 10% in Sub-Saharan Africa.[9] Moreover, <1% of women in four West African countries had ever been screened for cervical cancer.[10]

Although cervical cancer screening is proven to reduce cervical cancer incidence, many factors influence screening uptake.[11] The rate of screening uptake has been shown to vary by knowledge about cervical cancer and screening services, in addition to other factors, such as individual perception, beliefs, attitudes and culture and partner attitude.[12] Several studies have suggested that many women, particularly those with low levels of knowledge about cervical cancer and screening, may not recognise the benefit of screening over the possible consequences of forgoing screening.[13–18]

Although it is very limited in scope, there are prevention, treatment and rehabilitation strategies for cervical cancer, such as risk assessment, screening and clinical interventions, in Sub-Saharan Africa. Nevertheless, these services are not being fully used because of structural and behavioural barriers.[19,20] To enhance cervical cancer screening and treatment efforts, it is necessary to identify the factors influencing screening uptake in eligible women and their prevalence. Therefore, in this meta-analytic review, we aimed to estimate the pooled uptake of cervical cancer screening uptake and identify its predictors among Sub-Saharan African women.

Methods

The protocol has been registered with PROSPERO, an international prospective register of systematic reviews,[21] under registration number CRD42017079375. This meta-analytic review is reported in compliance with the recommendation of Preferred Reporting Items for Systematic Reviews and Meta-Analysis (PRISMA) 2015 statement.[22] The PRISMA Explanation and Elaboration document was followed and complemented by A Measurement Tool to Assess Systematic Reviews tool.[23] A PRISMA flow diagram[24] was used to illustrate the article screening and selection process (Fig. 1).

Literature search

PubMed, EMBASE, CINAHL, Web of Science, African Journals OnLine and Scopus electronic databases were explored to extract all available literature. Cross-references of included articles and grey literature were also searched. The search strategy (Table S1 in the supplementary material) was developed in consultation with medical information specialist and Peer Review of Electronic Search Strategies 2015 guideline statements.[25]

Eligibility criteria

Studies were included if they met the following inclusion criteria: (i) observational (i.e. cross-sectional, case-control, cohort) and (quasi) randomised controlled trial studies; (ii) conducted in Sub-Saharan Africa between January 2000 and December 2019; and (iii) published in the English language. Case reports, case series, expert opinions, qualitative studies, duplicated articles and studies with substantial incomplete data were excluded.

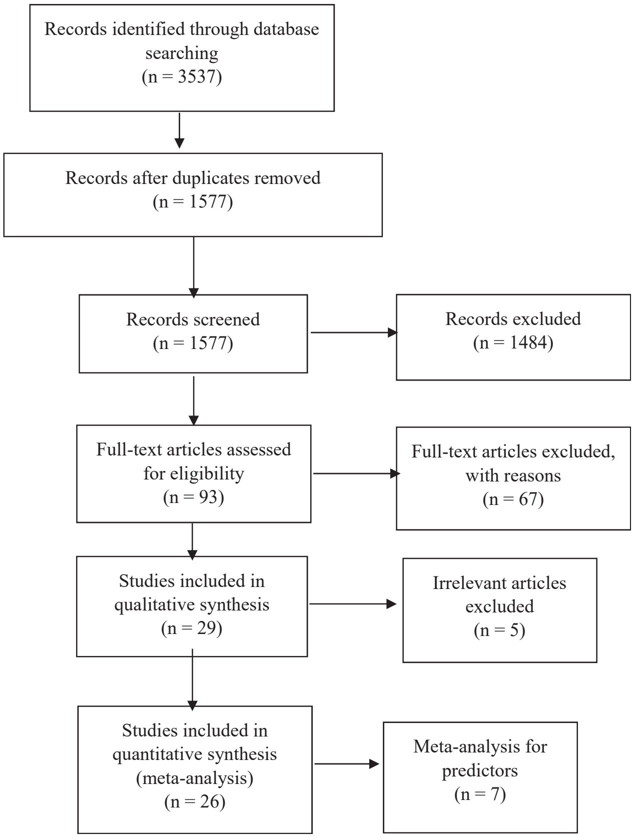

Fig. 1. PRISMA flow diagram for predictors of cervical cancer screening, January 2000 to January 2019. PRISMA, Preferred Reporting Items for Systematic Reviews and Meta-Analysis.

Study selection and data extraction

Initially, all identified articles were imported into Covidence.[26] After duplicate studies were excluded, two researchers (M.A.M. and N.B.Y.) identified articles by analysing the abstracts and titles for relevance to the proposed review topic. Agreement between the reviewers was made by consensus. Then, the full texts of the remaining articles were systematically reviewed for further eligibility. Finally, two reviewers (M.A.M. and N.B.Y.) extracted all relevant information, including first author, publication year, country, sample size, study design, prevalence,significant predictors and source of funding using an Excel spreadsheet. A disagreement between reviewers was solved through consensus.

Quality assessment of included studies

Two reviewers (N.B.Y. and M.A.M.) assessed the quality of selected articles using the Newcastle-Ottawa Scale (NOS) for cross-sectional studies.[27] The tool has three domains: selection (maximum of five stars), comparability (maximum of five stars) and outcome (maximum of five stars). In this review, studies were ranked as 'very good' if they scored ≥5 stars, 'good' for 4 stars, 'satisfactory' for 3 stars and 'unsatisfactory' for 0–2 stars. Quality assessment and funding sources of the studies are available in the supplementary material (Table S2).

Data analyses

An inverse variance-weighted random-effects model meta-analysis was performed to estimate the pooled uptake and odds ratio (OR) of predictors with 95% confidence interval (CI). To maintain adequate power, the meta-analysis was only used if at least five studies were available on a particular outcome of interest. Jackknife sensitivity analysis using the leave-one-out method was used to assess the effect of individual studies on the pooled OR estimate, significance level of estimate and between-study heterogeneity. The Jackknife is a linear approximation of the bootstrap, which systematically removes each observation of a data set, calculating estimates and finding the average of the calculation. A study was excluded when the pooled OR estimate increased or decreased by 1 and if there were changes to the significance level after removing that particular study from the meta-analysis. Owing to a small number of studies available for some variables, the change in heterogeneity threshold was not considered as a primary criterion to detect and exclude the outlier study. A narrative synthesis was used to summarise evidence on predictors. Heterogeneity between studies was tested using the Cochran's Q test and the Higgins's I^2 test statistic. The risk of publication bias was checked by visualising funnel plots and Egger's regression statistical tests. STATA, version 11 (StataCorp, College Station, TX, USA, 2009), was used for statistical analysis. To examine the source of heterogeneity, subgroup analysis was carried out based on sample size, the geographic distribution of the studies and year of publication.

Results

Characteristics of the studies

A total of 3537 studies were retrieved through database and manual searching. After removing duplicates (1577), 93 full-text articles were assessed for further eligibility. Finally, 29 articles with 36,374 women were included in the meta-analysis and qualitative analyses. Only seven studies were included in the meta-analysis for knowledge and cervical cancer screening (Fig. 1).

This review included studies conducted in the following Sub-Saharan African countries: 1 in Ghana, 1 in Burkina Faso, 1 in Botswana, 6 in Nigeria, 7 in Ethiopia, 4 in Kenya, 2 in Uganda, 2 in Tanzania, 2 in Zimbabwe, 1 in Mozambique, 1 in Cameroon and 1 in South Africa. All of the included investigations were cross-sectional studies. In total, 28 studies had a 'very good' quality score ($\geq$5 stars) and one study had 'good' quality score (4 stars) (Table 1).

Uptake of cervical cancer screening

The pooled uptake of cervical cancer screening in Sub-Saharan Africa was 12.87% (95% CI: 10.20, 15.54), and there was considerable heterogeneity ($I^2 = 98.5\%$). A random-effects model was used (Fig. 2), and subgroup analysis was conducted by region, sample size and year of publication. Based on the subgroup analysis, screening uptake ranged from 7.65% in the southern Sub-Saharan African countries to 14.13% in the eastern countries (refer to Fig. S1 in the supplementary material). By sample size, 13.83% of women were screened in a sample size group of <800, while 11.34% were screened in studies with sample sizes >800 (Figure S2). In addition, 13.5% of women were screened among studies published after 2015 (Figure S3). Sensitivity analysis was performed; no significant change was noted in the overall OR. There was publication bias, as evidenced by Egger's test P=0.048.

Predictors of cervical cancer screening

A study in Ghana and one in Ethiopia both showed that lack of formal education was significantly associated with low utilisation of cervical cancer screening services.[28,35] On the other hand, three studies[30,37,41] revealed that being HIV positive was a significant predictor for utilisation of the screening service. Awareness of place of screening also increased screening uptake in Kenya and Sudan.[37,45] An increase in cervical cancer screening was noted as age increases.[48] Tefera and Mitiku[34] reported a higher proportion of screened mothers aged 25–49 years. Similarly, Three studies[30,33,41] reported higher utilisation of the screening services with increasing age .

Negative attitudes, perceived susceptibility and perceived barriers have also been shown to reduce the likelihood of cervical cancer screening uptake.[41,42,55] Indeed, a positive attitude increased service utilisation in Ethiopia.[48] Akinyemiju et al.[44] in Nigeria reported that women were more likely to be screened if the provider was also female. On the contrary, not preferring gender of physician increased screening among Ethiopian women.[55] Two studies in Ethiopia reported that counselling about screening was associated with uptake of the service.[48,55] Abnormal vaginal bleeding,[28] heard about HPV and oral contraceptive use,[29] health insurance and condom use,[33] lack of awareness about the seriousness of cervical cancer,[36] fear of a bad result after screening,[37] multiple sexual partners and sexually transmitted diseases[41,48] and screening services provided at government health institutions[45] were also significantly associated with cervical cancer screening uptake (Table 1).

A meta-analysis of seven studies[29,30,33,34,39,41,55] revealed that knowledge about cervical cancer screening was significantly associated with cervical cancer screening (OR: 4.81; 95% CI: 3.07, 7.51). There was moderate heterogeneity ($I^2 = 47.8\%$); hence, a random effect model was used (Fig. 3). The Egger's test showed that no publication bias existed ($P = 0.44$).

Discussion

In this systematic review and meta-analysis, the overall uptake of cervical cancer screening was pooled from 26 studies in Sub-Saharan Africa and significant predictors of cervical cancer screening were identified. Knowledge about cervical cancer screening increased uptake of the service by nearly five times. In addition, educational level, age, HIV status, contraceptive use, perceived susceptibility and awareness about screening locations were predictors of cervical screening in Sub-Saharan Africa. The findings of this review revealed evidence to improve policies and practices aimed at addressing the utilisation of cervical cancer screening services across the region.

The pooled prevalence of cervical cancer screening in Sub-Saharan Africa was 12.12% (95% CI: 9.48, 14.76) in the present review. This rate is lower than that reported in studies of Chinese-Canadian and Malaysian women, which were 57%[56] and 48.9%,[57] respectively. Similarly, this rate is lower than that found in women with limited primary education in Indonesia (33–60%), Malaysia (23%) and Thailand (67.6%) but higher than that in women with limited primary education in the Philippines (7.7%) and Vietnam (4.9%).[58] However, these figures should be interpreted cautiously, as they are based on 2000–2001 World Health Organization estimates and may be out of date. Previous literature suggests that the lower uptake of screening in Sub-Saharan Africa may be due to overcrowding and overburden of healthcare

Table 1
Characteristics of the included studies from Sub-Saharan Africa, January 2000 to January 2019.

Study	Publication year	Country	Sample size	Screened women	Predictors	Quality score (stars)
Adanu et al.[28]	2010	Ghana	3183	25	Lack of formal education Abnormal vaginal bleeding	7
Sawadogo et al.[29]	2014	Burkina Faso	840	93	Heard about cervical cancer Knowledge about transmission mode Heard about human papillomavirus Oral contraceptive use	8
Mingo et al.[30]	2012	Botswana	376	271	Age 31–84 y Being HIV positive Heard about cervical cancer	7
Dim et al.[31]	2009	Nigeria	912	82	Not reported	5
Chigbu et al.[32]	2011	Nigeria	3712	389	Not reported	6
Cunningham et al.[33]	2015	Tanzania	575	35	Condom use Age 40–49 y, age >50 y Health insurance Knowledge about cervical cancer	7
Tefera and Mitiku[34]	2016	Ethiopia	634	68	Age 25–35 y, age 35–49 y Knowledge about cervical cancer	8
Aweke et al.[35]	2017	Ethiopia	595	58	Lack of formal education Primary education Secondary education	8
Morema et al.[36]	2014	Kenya	424	74	Lack of awareness about seriousness of disease	8
Orango'o et al.[37]	2016	Kenya	2505	273	Being HIV positive Fear of bad result Know place of screening	8
Tiruneh et al.[38]	2017	Kenya	9016	1750	Not reported	8
Lyimo and Beran.[39]	2012	Kenya	354	80	Knowledge about cervical cancer	8
Twinomujuni et al.[40]	2015	Tanzania	416	29	Not reported	8
Bayu et al.[41]	2016	Ethiopia	1286	235	Age 30–39 y Multiple sexual partners Sexually transmitted diseases Being HIV positive Knowledge about cervical cancer Perceived susceptibility and barriers	8
Idowu A et al.[42]	2016	Uganda	338	27	Negative attitude	8
Akanbi OA et al[43]	2015	Nigeria	737	110	Not reported	5
Akinyemiju et al.[44]	2015	Nigeria	1236	274	Female provider	8
Ahmed et al.[70]	2016	South Africa	500	79	Not reported	6
Ndejjo et al.[45]	2016	Uganda	845	43	Getting reproductive care at government facility Know place of screening Ease of getting reproductive service	8
Mupepi SC et al.[46]	2011	Zimbabwe	700	63	Knowledge of screening	8
Nwankwo et al.[47]	2011	Nigeria	845	36	Not reported	7
Bante et al.[48]	2019	Ethiopia	517	108	Age Counselling Positive attitude Visited health facility STIs	8
Brandão et al.[49]	2018	Mozambique	3177	96	Not reported	9
Donatus et al.[50]	2019	Cameroon	253	110	Not reported	4
Gebregziabher et al.[51]	2019	Ethiopia	344	59	Sexual experience Marital status Place of birth Year of study	7
Getachew et al.[52]	2019	Ethiopia	520	130	Not reported	8
Ifemelumma et al.[53]	2019	Nigeria	388	80	Not reported	6
Makurirofa et al.[54]	2019	Zimbabwe	409	15	Not reported	7
Nigussie et al.[55]	2019	Ethiopia	737	114	Government employee Know someone screened History of gynaecologic exam Gender of physician Counselling Knowledge Perceived susceptibility	8

STIs, Sexually Transmitted Infections.

providers at tertiary facilities.[59] Although cervical screening services are being offered, free of charge, in many African countries, out-of-pocket payment and fear of hidden charges were reported as barriers for utilisation of the service in some countries.[60] In addition, access to screening services, social support and other cultural and contextual factors might decrease utilisation of screening in Sub-Saharan Africa. As national screening campaigns have been promoted in recent years, the results from older studies might affect the pooled estimate of the present review. A root cause analysis in low-income countries reported that competing incentives among groups with shared interests in the service, suboptimal working conditions and lack of cervical cancer

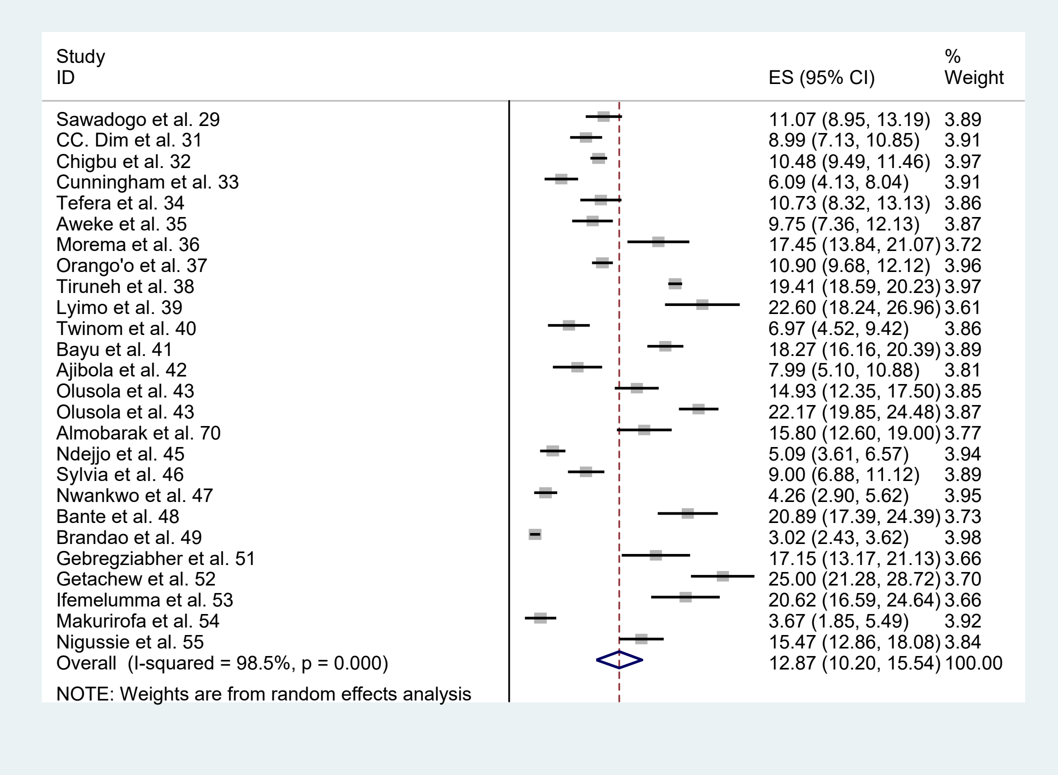

Fig. 2. Forest plot of pooled prevalence of cervical cancer screening in Sub-Saharan Africa, January 2000 to January 2019. CI, confidence interval; ES, effect size.

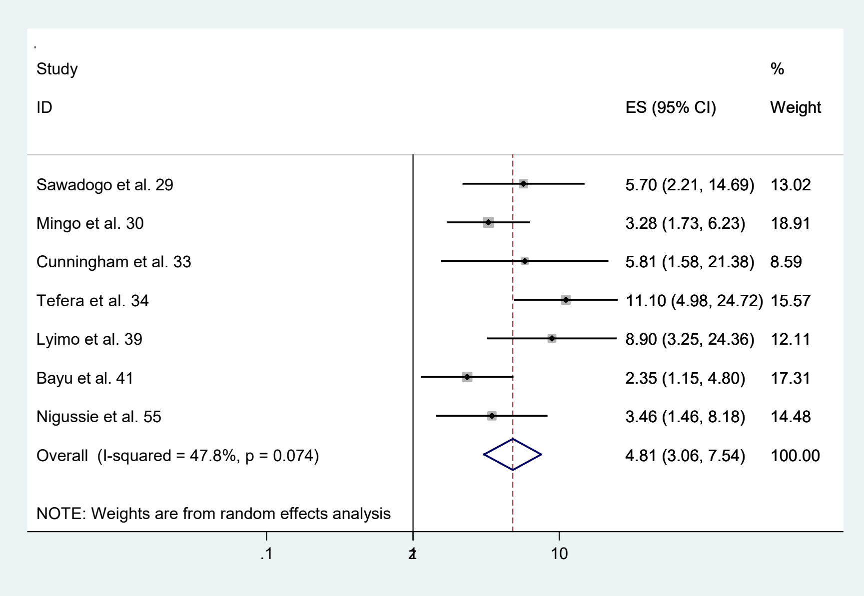

Fig. 3. Forest plot for knowledge about cervical cancer screening and uptake of service in Sub-Saharan Africa from January 2000 to January 2019. CI, confidence interval; ES, effect size.

prevention support in the political structures of the countries were identified as obstacles for successful cervical screening.[61] Another study, a Cochrane review of randomised trials, confirmed that invitations (e.g. appointments, letters, phone calls, verbal recommendations, prompts and follow-up letters) to women who were eligible for screening increased uptake of the service.[62] A systematic review in low- and middle-income countries revealed that telephone reminders or messages led to increasing cervical screening uptake.[63] Scaling up of screening services to all primary and secondary healthcare facilities and the use of trained medical staff may be important to increase uptake. Lower utilisation of screening services in Sub-Saharan Africa may also signal that political commitment is needed to improve cervical cancer prevention efforts.

The current systematic review revealed that lack of formal education and inadequate awareness about the seriousness of cervical cancer were associated with low utilisation of cervical cancer screening. This finding is consistent with a study in India that reported a higher incidence of cervical lesions among illiterate women due to their late presentation to health facilities.[59] Community mobilisation, including the use of village health promotors, may be important to increase uptake of screening services. In India, rural cancer registries and campaigns were found to be useful in detecting cervical cancer at the village level.[64] Moreover, the current review noted a higher utilisation of screening among older women, which is consistent with a study conducted in Malaysia.[57] This might be due to the fact that older women tend to seek treatment for their age- or hormone-related complaints. In the Netherlands, women aged 40–50 years who felt a high personal moral obligation had the highest likelihood of screening uptake.[65]

Women in the current review were more likely to have cervical cancer screening when the provider was female. Similarly, a study in Canada revealed that cervical cancer screening was associated with culturally sensitive healthcare services.[56] Together, these findings may imply the need for culturally appropriate care and outreach. Moreover, the current review showed that women tend to underuse the screening service owing to fear of bad results. Evidence shows that there are potential harms of screening, including anxiety related to positive results.[66] The present review also identified negative attitudes, perceived susceptibility and perceived barriers as significant factors for screening uptake. As women's beliefs may contribute to lower uptake of screening,[65] intervention strategies should focus on beliefs and attitudes about cervical cancer.

In the current review, women who knew about cervical cancer are nearly five times more likely to use cervical cancer screening than those who did not. Studies have shown that awareness about cervical cancer screening is a priority in resource-limited countries.[59] Similarly, general knowledge about cervical screening tests was associated with cervical cancer screening uptake among Chinese-Canadian women.[56] In addition, the current finding is in line with a study conducted in Malaysia[57] and systematic reviews in low- and middle-income countries.[67,68] Awareness about screening services might change the attitude of women to use the service. The role of community healthcare workers in educating the local population and raising awareness[69] needs to be highlighted and made a priority.

We registered our protocol prospectively, and reporting was based on established guidelines. We included all women who reported cervical cancer screening, regardless of screening modalities.

As a limitation, this finding might be prone to risk of bias due to the substantial heterogeneity of studies included from different locations. In addition, differences in cervical screening modalities across the included studies might influence the results of this review. In this review, only English language articles were included. Moreover, differences in how knowledge about cervical cancer was assessed in the included studies might affect the pooled estimates.

Conclusions

Cervical cancer screening uptake is low in Sub-Saharan Africa. Knowledge about cervical cancer was significantly associated with screening uptake. In addition, education level, age, awareness about screening locations, HIV status, attitude, provider gender, having heard about HPV, oral contraceptive use, health insurance, condom use, fear of a bad result, lack of awareness about the seriousness of the disease, multiple sexual partners, sexually transmitted diseases, counselling and receiving screening at public institutions were all important predictors of cervical cancer screening uptake in the region. Community-based education that is tailored to local culture, literacy level and pervasive attitudes is recommended to improve the uptake of cervical screening.

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Competing interests

None declared.

Author contributions

N.B.Y. and M.A..M. conducted the search and data extraction. N.B.Y., K.S. and M.T. wrote the first draft of the manuscript. N.B.Y., B.A. and H.K.M. conducted the statistical analyses. S.G., N.T.S. and T.D.H. contributed to data interpretation and the final editing of the manuscript.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2021.04.014.

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Short Communication

Children are safe in schools: a review of the Irish experience of reopening schools during the COVID-19 pandemic

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ABSTRACT

Objectives: Schools in the Republic of Ireland reopened to students and staff in late August 2020. We sought to determine the test positivity rate of close contacts of cases of coronavirus disease 2019 (COVID-19) in schools during the first half-term of the 2020/2021 academic year.

Methods: National-level data from the schools' testing pathway were interrogated to determine the positivity rate of close contacts of cases of COVID-19 in Irish primary, postprimary and special schools during the first half-term of 2020/2021 academic year. The positivity rates among adult and child close contacts were compared and the proportion of national cases of COVID-19 who were aged 4–18 years during the observation period was calculated to assess whether this proportion increased after schools reopened.

Results: Of all, 15,533 adult and child close contacts were tested for COVID-19 through the schools' testing pathway during the first half-term of the 2020/2021 academic year. Three hundred and ninety-nine close contacts tested positive, indicating a positivity rate of 2.6% (95% confidence interval: 2.3–2.8%). The positivity rates of child and adult close contacts were similarly low (2.6% vs 2.7%, $P = 0.7$). The proportion of all national cases of COVID-19 who were aged 4–18 years did not increase during the first half-term of the 2020/2021 school year.

Conclusions: The low positivity rate of close contacts of cases of COVID-19 in schools indicate that transmission of COVID-19 in Irish schools during the first half-term of the 2020/2021 academic year was low. These findings support policies to keep schools open during the pandemic.

As the coronavirus disease 2019 (COVID-19) pandemic has progressed, our understanding of the effects of the virus on children and their role in its transmission has increased. Schools are recognised as important places of education, safety, health-care and social learning.[1] The risks and benefits of schools remaining open must be balanced against the risks associated with COVID-19, and mitigated insofar as possible, in order to ensure the education of children in a safe setting.[1] Countries across the world have facilitated children returning to school at various stages of the pandemic. This return to education has been without evidence of increased rates of infection or transmission amongst school-age children, where appropriate preventative measures are in place.[2] The test positivity rate of close contacts of cases of COVID-19 in Irish primary, postprimary and special schools during the first half-term of 2020/2021 academic year in the Republic of Ireland was investigated by analysis of national-level school-specific data.

Infection prevention and control (IPC) measures in educational settings are essential to preventing and controlling COVID-19 transmission.[3] School-specific guidelines were developed for school staff and principals to support the safe reopening of schools.[4] Included in the guidance was advice on physical distancing, adapted based on age groups. Additional preventive measures included staggered class starting times, break and lunch times, assigning students to base classrooms with teachers transiting between classrooms and reconfiguration of classrooms using designated groupings of 'pods' and 'bubbles'. A 'bubble' is a class group which stays apart from other classes as much as possible. A 'pod' is a smaller grouping within the class 'bubble' with at least

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Table 1
Positivity rate among school close contacts (adults and children) by school type, 23 August 2020–28 October 2020, Republic of Ireland.

School type	No. schools	No. contacts	No. positive results (%, 95% CI)	No. negative results (%, 95% CI)	Unknown results (%, 95% CI)
Primary	372	9858	270 (2.7%, 2.4–3.1%)	9571 (97.1%, 96.7–97.4%)	17 (0.2%, 0.1–0.3%)
Postprimary	199	4897	101 (2.1%, 1.7–2.5%)	4785 (97.7%, 97.3–98.1%)	11 (0.2%, 0.1–0.4%)
Special	33	778	28 (3.6%, 2.4–5.2%)	749 (96.3%, 94.7–97.5%)	1 (0.1%, 0.001–0.7%)
Total	**604**	**15,533**	**399 (2.6%, 2.3–2.8%)**	**15,105 (97.2%, 97.0–97.5%)**	**29 (0.2%, 0.1–0.3%)**

CI, confidence interval.
Source: Schools' COVID-19 testing pathway.

1 m physical distance between individual pods. In addition, all staff and postprimary school students were requested to wear a face covering when a physical distance of 2 m could not be maintained. Enhanced cleaning of commonly touched surfaces, increased ventilation and adherence to rigorous respiratory and hand hygiene was also advised.[4]

Approximately one million staff and students attend school on a daily basis in the Republic of Ireland. When a case of COVID-19 in a school is notified to a regional Department of Public Health (DPH), the school is contacted and a public health risk assessment is performed. This assessment involves collecting and consideration of information from the case interview, local community outbreaks, and particularly information from the school on adherence to the previously outlined guidance. The standardised close contact definition is then applied to identify close contacts, and to determine the control actions required.

In Ireland, there is a bespoke testing pathway for close contacts identified in schools. Positive cases of COVID-19 identified through this schools' testing pathway are notified to the regional DPHs and case data are aggregated on the national Computerized Infectious Disease Reporting (CIDR) system.[5] Notification completeness is optimised by systematised electronic reporting from the laboratories. The data presented were extracted from two sources: total numbers of cases aged 4–18 years were obtained from CIDR and the positivity rate among close contacts in schools was calculated through data collected from the schools' testing pathway.

From August 23rd 2020–October 28th 2020, close contacts (adults and children) from 604 schools were referred for COVID-19 testing through the schools' testing pathway. Of these schools, 372 (61.6%) were primary, 199 (32.9%) were postprimary and 33 (5.5%) were special schools. In total, 15,533 individuals were identified as close contacts in the school setting and referred for COVID-19 testing (Table 1). This represents approximately 1.5% of the national school community. Of these, 399 close contacts (2.6%, 95% CI: 2.3–2.8%) tested positive and were spread across 156 schools. The highest positivity rate occurred in special schools (3.6%).

Of the 15,533 close contacts, 13,408 were children and 2125 were adults. The positivity rates of child and adult close contacts were similarly low at 2.6% and 2.7%, respectively ($P = 0.7$). Children in special schools had the highest positivity rate (4.4%) compared with children in primary (2.7%) and postprimary schools (2.1%). Adults in primary schools had the highest positivity rate (3.0%) compared with adults in postprimary (2.1%) and special schools (2.9%).

During the observation period, the proportion of COVID-19 cases nationally who were of school-going age (i.e. 4 18 years) did not increase and represented 14.3%, 15.0% and 14.9% of national cases during the months of August, September and October, respectively ($P = 0.6$). Of the 604 schools where close contacts were referred for testing, one school was advised by a regional DPH to close to reduce the risk of further COVID-19 transmission in the school.

This paper presents the transmission dynamics of COVID-19 in Irish primary, postprimary and special schools over a nine-week time period from when schools reopened in late August until the midterm break in late October 2020. Overall, the positivity rate of close contacts across all school types in the Republic of Ireland during this period remained at a stable, low level. The positivity rate of close contacts in schools was considerably lower than the positivity rate of close contacts in the community at the time, which was approximately 10%. This finding of a low positivity rate of close contacts of cases of COVID-19 in schools is consistent with findings from similar studies conducted nationally[6] and internationally[2] and supports the observation that children are not the main drivers of COVID-19 transmission.[7] While it is a proxy measure, the low positivity rate of close contacts in schools suggests that the level of intraschool transmission of COVID-19 which occurred in the Republic of Ireland during this time period was low. Additional data from regional DPHs on individual schools, where in-depth epidemiological investigations were conducted, supported the fact that intraschool transmission of COVID-19 was uncommon during the nine-week observation period.[8] Many close contacts who subsequently became cases were reported to have had other, more likely, sources and exposures for COVID-19 infection. The reporting of a low positivity rate among school close contacts is further supported by the observation that the proportion of 4–18-year-old cases among all national cases did not increase when schools in the Republic of Ireland reopened in late summer. The 14-day national incidence of COVID-19 in the Republic of Ireland increased from 33 cases per 100,000 in early September 2020 to 299 cases per 100,000 in late October 2020,[9] but the proportion of cases who were school-going aged children remained remarkably stable.

The positivity rates amongst adults and children in schools were both low. This finding is somewhat surprising as the available evidence to date suggests children may be less susceptible to infection with COVID-19 than adults.[10] The reasons for this result, however, may be due to IPC measures implemented in schools which mitigated against virus transmission in all settings and in all age groups.

Prolonged school closures have been linked to detrimental effects on many aspects of children's health and well-being.[1] This interrogation of national-level Irish data investigating the transmission dynamics of COVID-19 contributes to the emerging evidence and demonstrates that, with appropriate mitigation measures, the rate of onward transmission of COVID-19 in educational facilities remains low. Furthermore, it supports policies on schools remaining open for the educational and psychosocial development of children.

In light of the Irish experience thus far, we recommend that the facilitation of onsite learning in all school types, with all mitigation measures carefully adhered to, remains a high priority. Continued evaluation of the pandemic, timely case investigation, and contact tracing and testing should be prioritised to ensure close contact positivity remains low among staff and students and onsite learning can be safely maintained.

Author statements

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Ethical approval

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None declared.

Authorship

AC collated the data and produced an initial report. PW conducted further analysis on the data. PW, EK, RC, AC, MW and MOS wrote and edited the article for submission.

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Short Communication

Evaluation of a measles vaccination campaign at the universities in the city of Zurich, 2019

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ABSTRACT

Objectives: In 2019, there were 29 reported cases of measles in the Canton of Zurich, with two cases occurring among university students. In collaboration with the University of Zurich Travel Clinic, the Health Department of the Canton of Zurich offered free measles vaccination to all employees and students at the University of Zurich (UZH) and the Swiss Federal Institute of Technology (ETH). This short communication shares the results of this large measles vaccination campaign.
Study design: Vaccination intervention campaign.
Methods: All employees and students at the UZH and ETH were informed via an email distribution list that they were eligible for cost-free consultation and measles vaccination (when indicated). Consultations and immunizations took place over the course of 3 days in June 2019 at the UZH Travel Clinic. All those who were missing one or two doses of measles vaccination, and had no contraindications, were vaccinated. Booster immunizations were offered until December 2019.
Results: A total of 411 individuals participated in the campaign. Thirty-five individuals (8.5%) were found to have sufficient measles vaccination on consultation and received no additional vaccination. A total of 376 individuals (91.5%) met the eligibility criteria and were vaccinated; 83 individuals (20.2% of all participants and 22.1% of those vaccinated) returned for a second vaccination. In total, the campaign saw 494 visits (including consultations without immunization and visits for second immunization). Demographic data were collected for 439 visits where measles vaccination was administered. From these, 51.7% were for an individual's first measles vaccine dose, 27.3% for a second dose, 18.9% for a booster immunization and 2.1% were unknown. 54.7% of campaign visits were made by females; and 45.0% of visits were made by those aged 18–29 years, 27.9% by those 30–39 years, 14.6% by those 40–49 years, and 12.6% by those 50+ years. 49.8% of visits were made by students and 48.5% by employees. More students needed the first dose (54.2% of first-dose visits), whereas more employees received booster immunization (57.8% of booster visits).
Conclusions: The measles vaccination campaign was well attended, particularly by the younger age group 18–29 years and females. Coupled with intense media attention, such a campaign immediately following an outbreak may be an effective method to increase vaccination coverage.

Introduction

Measles is widely accepted as a highly contagious and potentially dangerous disease. In 2018 alone, measles accounted for more than 142,000 deaths worldwide.[1] This is particularly unfortunate as the disease can be easily prevented by vaccination. In line with this, the WHO estimates that measles vaccination prevented more than 23 million deaths between 2000 and 2018.[1] In recent years, there has been a surge in measles outbreaks. In Europe, more than 100,000

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measles cases were reported between January and October 2019; this is more than those reported in all of 2018 and over three times the total reported in 2017.[2] Significantly, the majority of cases occur among unvaccinated individuals, further underscoring the importance of measles immunization.

As with other European countries, the campaign to eliminate measles from Switzerland was recently extended by the Swiss Federal Office of Public Health to 2020. It is unlikely that this target will be met, although vaccination coverage has increased considerably. Compared with the national measles vaccination coverage of ca. 73% for children in 2007, data from 2015 to 2019 showed coverage estimates have risen to 90%, 94%, and 94% at two doses for children of age 2, 8 and 16 years, respectively, and 88% for adults 20–29 years of age.[3,4] Despite this high coverage, and despite having been declared ‘measles endemic free’,[5] multiple measles outbreaks have been reported, with cases predominantly in older children and young adults, similar to other high-income, high vaccination coverage countries.[5,6]

Although much smaller than the nationwide outbreak between 2006 and 2009 where 4415 measles cases were reported (incidence 58/100,00), there were several measles outbreaks in Switzerland in 2019, with 221 reported cases (incidence 2.57/100,000), including two deaths.[7,8] In the Canton of Zurich alone, 29 cases (incidence 1.91/100,000) were reported, with two occurring in university students[7] (personal communication). As a preventive measure, the Health Department of the Canton of Zurich conducted a vaccination campaign offering free consultations and measles vaccination (where indicated) to all students and employees of the University of Zurich (UZH) and the Swiss Federal Institute of Technology (ETH). Consultations/vaccinations took place at the UZH Travel Clinic (TravClin). This short communication describes the reach of this vaccination campaign with information about the visits and vaccinations administered.

Measles vaccination campaign

All employees and students of the UZH and ETH were informed via an email distribution list that they could receive a cost-free consultation and, if indicated, measles vaccination on 3 days in June 2019 at the TravClin. Priorix, the combined measles, mumps, and rubella (MMR) vaccine, and Measles Live Vaccine, which only protects against measles were offered. Additional immunizations (second vaccination or booster vaccination, if indicated) were administered to participants through December 2019. Booster vaccinations were defined as having both measles vaccine shots administered during the campaign. All participants were requested to bring their vaccination card; those without a vaccination card were requested to provide verbal confirmation from their parents and/or healthcare provider regarding their vaccination status.

A total of 411 individuals participated in the campaign. Thirty-five individuals (8.5%) were found to have sufficient measles vaccination on consultation and were not vaccinated. A total of 376 individuals (91.5%) were vaccinated; 83 individuals (20.2% of all participants and 22.1% of those vaccinated) returned for an

Table 1

Characteristics of the visits in the measles vaccination campaign in June–December 2019.

Characteristics	n	Vaccinations					
		Visits	Total	MCV first dose	MCV second dose	Booster	Undeclared
		480 (%)*	439 (%)*	227 (%)*	120 (%)*	83 (%)	9 (%)*
Demographics							
University							
ETH		266 (55.4)	240 (54.7)	128 (56.4)	68 (56.7)	39 (47.0)	5 (55.6)
UZH		211 (44.0)	197 (44.9)	99 (43.6)	51 (42.5)	44 (53.0)	3 (33.3)
Undeclared		3 (0.6)	2 (0.5)	0 (0.0)	1 (0.8)	0 (0.0)	1 (11.1)
Role							
Students		239 (49.8)	219 (49.9)	123 (54.2)	58 (48.3)	34 (41.0)	4 (44.4)
Employees		233 (48.5)	214 (48.7)	101 (44.5)	61 (50.8)	48 (57.8)	4 (44.4)
Both		5 (1.0)	4 (0.9)	3 (1.3)	0 (0.0)	1 (1.2)	0 (0.0)
Undeclared		3 (0.6)	2 (0.5)	0 (0.0)	1 (0.8)	0 (0.0)	1 (11.1)
Gender							
Female#		212 (55.1)	193 (54.7)	90 (54.5)	58 (57.4)	44 (53.0)	1 (25.0)
Male#		173 (44.9)	160 (45.3)	75 (45.5)	43 (42.6)	39 (47.0)	3 (75.0)
Age group							
18–29		215 (44.8)	197 (45.0)	106 (46.7)	45 (37.5)	41 (49.4)	5 (55.6)
30–39		132 (27.5)	122 (27.9)	62 (27.3)	41 (34.2)	16 (19.3)	3 (33.3)
40–49		73 (15.2)	64 (14.6)	28 (12.3)	24 (20.0)	11 (13.3)	1 (11.1)
50–55		32 (6.7)	31 (7.1)	16 (7.0)	7 (5.8)	8 (9.6)	0 (0.0)
56+		27 (5.6)	24 (5.5)	14 (6.2)	3 (2.5)	7 (8.4)	0 (0.0)
Undeclared		1 (0.2)	1 (0.2)	1 (2.4)	0 (0.0)	0 (0.0)	0 (0.0)
Time and type of vaccination							
June		350 (72.9)	310 (70.6)	209 (92.1)	93 (77.5)	0 (0.0)	8 (88.9)
Booster		83 (17.3)	83 (18.9)	0 (0.0)	0 (0.0)	83 (100.0)	0 (0.0)
New vaccination mid Jun–Dec		47 (9.8)	46 (10.5)	18 (7.9)	77 (22.5)	0 (0.0)	1 (11.1)
MCV first dose			227 (51.7)				
MCV second dose			120 (27.3)				
MCV booster			83 (18.9)				
Undeclared			9 (2.1)				
Priorix			436 (99.3)				
Measles Live Vaccine			3 (0.70)				

Free measles vaccinations were offered on June 3, 5, and 12, 2019 at the Zurich Travel Clinic. Boosters were offered between mid-June and December 2019. MCV, measles containing vaccine. Either Priorix (measles, mumps, rubella, MMR) or Measles Live Vaccine was administered. ETH: Swiss Federal Institute of Technology; UZH: University of Zurich. * 494 visits were recorded, but data were collected for only 480 visits. #n = 385 (gender was not collected on the first day of the campaign).

additional (second or booster) vaccination. In total, the campaign saw 494 visits (including consultations without immunization, initial immunizations, and visits for repeated immunizations by the same participant). Demographic data were collected for 439 visits where measles vaccination was administered. In 56.0% of cases, participants brought a vaccination card. From these 439 visits, 51.7% were for an individual's first known/recorded measles vaccine dose, 27.3% for a second known/recorded dose, and 18.9% for a known/recorded booster immunization. 2.1% were unknown. 54.7% of campaign visits were made by females and 45.3% by males. 45.0% of visits were made by individuals aged 18–29 years, 27.9% by those 30–39 years, 14.6% by those 40–49 years, and 12.6% by those 50+ years. 49.8% of visits were made by students and 48.5% by employees. More students than employees received their first known/recorded dose (54.2% vs 44.5% of first-dose visits), whereas more employees than students received a known/recorded booster immunization (57.8% vs 41.0% of booster visits). 99.3% of immunizations were performed with Priorix and 0.7% with Measles Live Vaccine (see Table 1).

Lessons learned

In 2014–2016, prior to the vaccination campaign described here, the Health Department of the Canton of Zurich sponsored a similar measles vaccination campaign, but on-site at the UZH campus for all students and employees of the UZH and ETH in collaboration with the UZH Travel Clinic. During the 2014–2016 campaign, however, only 212 consultations with 148 vaccinations were performed. Similarly, in 2017, an on-site measles vaccination campaign was organized at four trade schools in the city of Zurich, where only 17 individuals visited, of which 10 were vaccinated.

Significantly, preceding our 2019 campaign, where 376 individuals were vaccinated at a single site within 3 days, there had been extensive media coverage of the national measles outbreaks, as well as outbreaks abroad.[9,10] Furthermore, both the UZH and ETH released multiple communications to students and staff, particularly in regard to the 2 documented cases among university students. A quarantine period of 3 weeks was imposed for all exposed, unvaccinated individuals. It is likely that these factors contributed to the increased attendance of this vaccination campaign. To increase vaccination uptake, a monovalent measles vaccine was also offered during the campaign to reach those who might refuse the traditional combined MMR vaccine; however, only 0.7% received this monovalent vaccine, indicating that the majority are not concerned with the MMR vaccine.

As the UZH Travel Clinic is a well-known vaccination center with the necessary infrastructure and trained personnel readily available and part of the University, it was possible to quickly organize a large vaccination campaign in response to the two student measles cases. In addition, the use of existing infrastructure helped to maintain a relatively low cost. The total cost of our 2019 vaccination campaign was 30,000 CHF ($31,000), or about 80 CHF ($83) per person. A recent review of 11 measles outbreaks from the United States between 2001 and 2018 estimated that the median cost of a single measles case was about $32,800, with a median cost of $223 per contact.[11] This suggests that a well-timed and well-executed campaign can be effective as well as cost-efficient. This experience could be useful especially in planning future vaccination campaigns, particularly in outbreak situations.

Conclusions

This university-based measles vaccination campaign was well attended, particularly by individuals 18–29 years. The timing of the campaign (in the context of an outbreak situation) and the resulting media attention likely helped to increase participation and, therefore, vaccine coverage. Additionally, the use of existing infrastructure greatly facilitated the speed at which the campaign could be implemented and helped to reduce costs. As it is important not to miss opportunities to promote vaccination, a strategy should be in place to quickly organize such campaigns in outbreak situations.

Author statements

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Original Research

Linking national public services data to estimate the prevalence of intellectual disabilities in The Netherlands: results from an explorative population-based study

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ABSTRACT

Objectives: Individuals with intellectual disabilities (ID) depend on public services for daily support and medical care; however, this group of individuals can be difficult to identify within population data. This therefore limits the opportunities to accurately estimate the size of the population with ID, monitor trends and tailor public health interventions according to the needs and characteristics of this group. By linking relevant databases, this study sought to identify individuals with ID in national data, to estimate the prevalence of ID based on public service use and to explore how this method can be used to better monitor the population with ID.

Study design: Explorative data linkage study using the Dutch population register and databases from public services accessible with an ID diagnosis.

Methods: The overall prevalence of ID in the Dutch adult population was estimated, specified by age group and sex, and the identified ID groups were also characterised by their support needs. Participants included the entire adult Dutch population who were alive on 1 January 2015.

Results: After linking databases, 187,149 adults with ID were identified within a population of almost 12.7 million Dutch adults, giving an ID prevalence estimate of 1.45%. Prevalence of ID was higher among males (1.7%) than females (1.2%). Most individuals with ID were identified through the use of residential care services (n = 91,064; 0.7%). Non-residential ID-related care was used by 27,007 individuals (0.2%). Social, employment or income support due to a (mild) ID was received by 69,078 individuals (0.5%); the mean age in these ID groups was between 8 and 10 years which is younger than that in the general Dutch population. ID prevalence declined with increasing age across all ID subgroups.

Conclusions: The ID prevalence in The Netherlands, as determined by ID-related public service usage, aligns with international estimates. This suggests that national supportive services are accessible and used by individuals with ID. Moreover, this demonstrated that databases from national supportive services can be a useful resource to identify individuals with ID at the population level and can enable structural monitoring of the ID population through linking national databases.

Introduction

Within every population, individuals with intellectual disabilities (ID) form a specific subgroup with support needs for health, behavioural and social problems. In many countries, including The Netherlands, these needs are met through a variety of public health and social support systems.[1–3] Many aspects related to health and well-being, including the care and services provided through these systems, are registered as part of routine practice, either automatically or manually.[4] These administrative data thus contain a wealth of information that could inform policy and practice about the characteristics and changes in the needs of those requesting and using these services.[5–7] However, in many countries, it is not apparent how individuals with ID can be identified within these data or how the data can be used for monitoring purposes.[8,9]

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For ID service providers, the infrastructure to set up data-linkages is also limited. Service providers can each have their own definition of ID, which limits opportunities to identify individuals across different databases, especially as many countries rely on multiple systems without national coverage.[10–14] Consequently, inaccurate or incomplete identification of individuals with ID may result in findings that are neither representative nor generalisable. Such findings subsequently have limited relevance for use in policy and practice. While administrative population data have the potential to accurately identify everyone with ID, a recent review showed that ID prevalence estimates through this method still vary substantially.[14]

In The Netherlands, many health and social systems have national coverage and uniform procedures to collect and combine information. Regardless of which definition for ID was used to enter the supportive systems (e.g., International Classification of Diseases [ICD]-10 or Diagnostic and Statistical Manual of Mental Disorders [DSM]-V), service use and support needs are categorised the same for all individuals using these systems. Furthermore, one would expect a correlation between ID severity and the level of support requested. This allows for reproducible and consistent methods to identify the Dutch population with ID, link databases and monitor developments at the population level. However, to date, the use of these databases has been limited. To allow better use of these existing resources, this study describes a method that uniquely identifies individuals with ID by linking national databases and provides an ID prevalence estimate based on this method.

Methods

Data sources and setting

This cross-sectional study used non-public microdata, which, under certain conditions, are accessible for statistical and scientific research from Statistics Netherlands, the Dutch national statistics office.[c] We took 2015 as the reference year and linked the Dutch population registry with two databases; one for chronic care and one for welfare support. Both of these databases contain information on utilisation of services accessible to individuals with ID who had formally been diagnosed by a healthcare professional according to their respective guidelines and diagnostic systems (e.g., ICD-10 or DSM-V).

The chronic care database contained information on all long-term care provided under the Chronic Care Act (CCA). The CCA regulates residential and other long-term care for people with chronic somatic conditions, vulnerable elderly people, people diagnosed with physical and/or (mild) ID or chronic mental illnesses.[15] Support through the CCA can either be in-kind by predefined care packages for residential care (ranging from 1 to 8 depending on the level of support needs) or self-arranged for people without residential care needs.

The welfare database included information on individuals active in sheltered workplaces, recipients of unemployment or disability benefits and the reason why the provision was called upon, which could include the presence of a mild ID diagnosis. The reason Mild-ID is labelled separately in these (income-related) systems is because the definition for Mild-ID (DSM-V/ICD-10) specifies “many adults will be able to work”.[16,17] Therefore a Mild-ID diagnosis grants access to these welfare systems, while a more severe ID diagnoses do not.

As additional regulations apply to children aged <18 years and most of the welfare services only concerned for individuals aged ≥18 years, we restricted this study to the Dutch adult population.

The study protocol for this exploratory study was reviewed by the Radboud University Medical Center institutional Ethics Committee who passed a positive judgment and waived the need for formal ethical assessment (2017–3921). We followed the Strobe checklist for cross-sectional studies[18] (see supplemental materials A).

Procedures for linkage and establishing ID groups

Before database linking was commenced, any multiple registrations of individuals within the same database were removed. Under the CCA, individuals can receive multiple entitlements within one year, resulting in multiple registrations, but only one can be active at a specific time. We therefore selected the most recent CCA registration per individual in 2015, under the assumption that it would reflect the most up-to-date support care needs. In the welfare database, individuals could simultaneously be registered as a recipient of multiple benefits, as one type of benefit could supplement another. Here, we recoded multiple records per individual into a single variable representing the combination of benefits received.

Linkage then started by retrieving sex, date of birth and a unique personal identifier (RIN number) as the primary variables from the population register containing all Dutch adults who were alive on 1 January 2015. Based on matching RIN numbers, we added information, if any, from the CCA database on entitlements to chronic care services, the type of care they were receiving and both the primary and secondary reason for which the CCA was called upon (A maximum of two reasons can be given.). Databases from the welfare systems were also linked by RIN number, which provided information about individuals entitled to benefits due to a mild ID diagnosis.

In the newly composed data set, we grouped individuals according to their support needs and degree of independence. We grouped all individuals receiving residential ID care through the CCA, regardless of any other registration, as having the greatest support needs and being least independent (residential ID group). All other individuals who called upon the CCA for any other ID-related reason and did not receive residential care were seen as having moderate support needs and being moderately independent (non-residential ID group). All individuals who had a Mild-ID diagnosis noted as their reason for accessing services in any of the databases were assumed to be the most independent ID group with the least support needs (Mild-ID). Through this classification, all individuals identified with an ID could be uniquely assigned to one of the three ID groups (i.e., residential ID group, non-residential ID group or Mild-ID). All remaining individuals who were not assigned to one of the three ID groups were assumed to be a member of the general population. A flow chart is presented in Fig. 1, and the classification rules used to define groups are specified in Supplemental materials B.

Statistical analyses

Demographics were presented as frequencies with percentages or means with standard deviation (SD). Age was grouped mid-decade to mid-decade (e.g., 35–44 years) following epidemiological conventions.[19] ID prevalence was calculated as the number of individuals identified with ID divided by the total number of people enrolled in this study. We specified frequencies by sex and 10-year age groups and provided separate prevalence estimates per ID subgroup. As we used population data, all prevalence estimates yielded very small confidence intervals, which were considered

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[c] Procedures can be found at: https://www.cbs.nl/en-gb/our-services/customised-services-microdata/microdata-conducting-your-own-research, or for further information. microdata@cbs.nl.

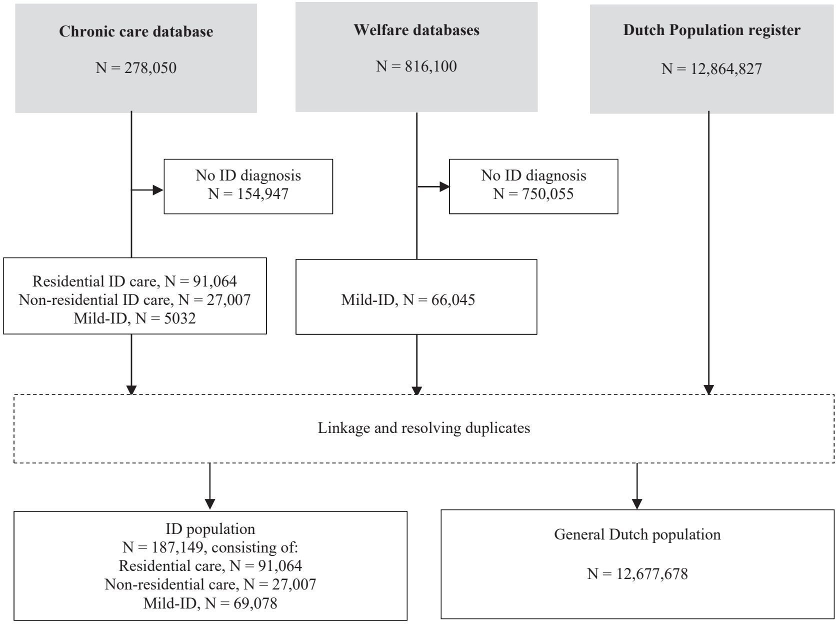

Fig. 1. Flow chart of study population. ID, intellectual disabilities.

uninformative and hence not included in the tables. Per subgroup, we used the available information to further characterise each ID subgroup separately. Analyses were conducted using SPSS (version 25.0).

Results

After linking the three databases (i.e., Dutch population registry, chronic care and welfare support databases), 187,149 individuals were identified with ID and could be assigned to one of the three ID groups based on their service use. The remaining 12,677,768 individuals were without any record of supportive ID services and were believed to form the general Dutch adult population (GenPop). Combined, the three ID groups gave an ID prevalence of 1.45% (95% confidence interval, 1.44–1.46). Mean ages across the ID groups were between 7.9 and 9.9 years which are lower (M_{age} range, 38.4–40.4 years) than those in the GenPop (M_{age}, 48.3 years). The residential care (56.0% males) and Mild-ID (63.4% males) groups contained more males than the GenPop (48.9%; Table 1). The overall ID prevalence was higher among males than females across all age groups and declined with increasing age, ranging from 3.6% for males between 18 and 24 years of age to 0.2% among females aged ≥75 years (Fig. 2).

Residential care

The residential care group consisted of 91,064 individuals with ID (56.0% male), with an average age of 40.3 years (SD = 16.4). The largest age group was the 18–24 years group (23.9%; Table 1). Almost half of the individuals in this group (n = 42,391, 46.5%) were supported by care package 3 (24.5%) or care package 6 (22.0%) and thus required “medium” to “intensive” support in daily living, had medium care needs and were “limited” to “very limited” in their independence (Table 2). The complete overview of distribution across care packages for residential care is presented in Table 2.

Non-residential care

The non-residential care group consisted of 27,007 individuals with ID (46.7% male), with an average age of 40.4 years (SD = 15.0). The largest age group was the 25–34 years group (22.8%; Table 1). The majority of individuals in this group (n = 20,955, 77.6%) self-arranged their supportive care without further specification in the CCA database. Among those whose care utilisation was specified, 2149 individuals (8.0%) received outpatient treatment and care, 959 individuals (3.6%) received generic, not ID-specific care, and the remaining 2944 individuals (10.9%) received care for another primary reason other than ID (e.g., for a physical disability or mental health). Having another reason to call upon the CCA besides an ID was common in this subgroup. In the group of individuals with ID who had more than one underlying reason for using care (n = 14,100, 52.2%), the ID was the primary reason for 8758 individuals, and for 5342 individuals, the ID was a secondary reason. Most prevalent conditions besides an ID were a psychiatric condition (n = 7433, 27.5%) or a somatic condition (n = 3494, 12.9%). Further characteristics of this group are shown in Table 3.

Mild-ID

The Mild-ID group consisted of 69,078 individuals (63.4% male), with an average age of 38.4 years (SD = 14.6). The largest age group was the 18–24 years group (25.9%), and 1828 individuals (2.7%) were aged ≥65 years (Table 1). Most individuals were supported by a single service only, being related to work in sheltered workplaces (n = 49,861, 72.2%), receipt of disability benefits (n = 11,631, 16.8%) or chronic care (n = 3033, 4.4%). The remaining 4553 individuals (6.6%) were supported through a combination of two or more of these services (Table 4).

Discussion

By linking three national databases, this is the first study to have identified individuals with ID in administrative population data in

Table 1
Demographics of the general adult Dutch population (GenPop) and the ID subgroups[a].

Demographics	GenPop, N = 12,677,768 (98.5%)		ID groups combined N = 187,152 (1.45%)					
			Residential care, N = 91,064 (0.7%)		Non-residential care, N = 27,007 (0.2%)		Mild-ID, N = 69,078 (0.5%)	
	N	*%*	*N*	*%*	*N*	*%*	*N*	*%*
Sex								
Male	6,196,789	48.9	50,983	56.0	12,624	46.7	43,763	63.4
Female	6,480,979	51.1	40,081	44.0	14,383	53.3	25,315	36.6
Age, M (SD)	48.3 (17.8)		40.3 (16.4)		40.4 (15.0)		38.4 (14.6)	
18–24 years	1,362,047	10.7	21,727	23.9	4727	17.5	17,892	25.9
25–34 years	1,931,948	15.2	17,898	19.7	6160	22.8	12,557	18.2
35–44 years	2,225,395	17.6	14,626	16.1	5739	21.3	12,149	17.6
45–54 years	2,457,868	19.4	16,584	18.2	5421	20.1	15,104	21.9
55–64 years	2,104,631	16.6	12,519	13.7	3143	11.6	9478	13.7
65–74 years	1,551,963	12.2	5640	6.2	1260	4.7	1870	2.7
≥75 years	1,043,916	8.2	2070	2.3	560	2.1	28	0.04

ID, intellectual disabilities.

[a] Total population size to base prevalence calculation on was n = 12,864,827.

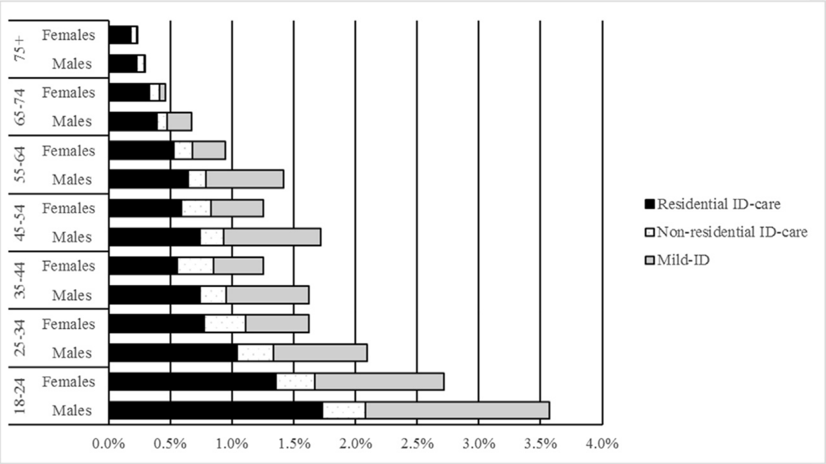

Fig. 2. Prevalence estimates per ID severity group, by age and sex. ID, intellectual disabilities.

The Netherlands and to consequently provide an estimate for the ID prevalence in the Dutch adult population. Combining administrative data on chronic care and welfare benefits gave an ID prevalence estimate of 1.45%. Based on the level of support needs, the ID population could be classified as residential care, non-residential care or Mild-ID, with a prevalence of 0.7%, 0.2% and 0.5%, respectively.

Good quality data should be the basis for decision-making in the policy and practice of public health. However, access to accurate data has been lacking in the ID field, especially at the population

Table 2
Distribution of individuals with ID across residential care packages[b].

Package	Level of required guidance in daily living and activities (0–4)[a]	Level of required care (0–4)	Level of Independence (0–4)[a]	Total, n (%)
1	1-Some	0-None	1-Fairly independent	734 (0.8)
2	2-Medium	0-None	2-Some limitations	4639 (5.1)
3	2-Medium	2-Medium	2-Some limitations	22,350 (24.5)
4	2-Medium	3-Intensive	3-Very limited	11,949 (13.1)
5	3-Intensive	3-Intensive	3-Very limited	10,733 (11.8)
6	3-Intensive	3-Medium with behaviour therapy	3-Very limited	20,041 (22.0)
7	4-Very intense	3-Medium with behaviour therapy	3-Very limited	12,867 (14.1)
8	2-Medium	4-Complete care and nursing	4-Completely dependent	7751 (8.5)
				91,064

ID, intellectual disabilities.

Source: User guide Intellectual disabilities—Information per care package, National health care institute, and Care profiles, Care Needs Assessment centre (CIZ), https://www.ciz.nl/images/pdf/beleidsregels/Zorgprofielen.pdf.

[a] Categories 0-No guidance required and 0-Completely independent do not occur.

[b] Package refers to the pre-arranged care packages available to individuals in this group.

Table 3
Distribution of individuals with ID across types of non-residential care.

Description of care	ID primary diagnosis	ID secondary diagnosis	Total, n (%)
Self-arranged without residency, not specified	19,217	1738	20,955 (77.6)
Functional treatment without admission or residency	2179	7	2186 (8.1)
Sensory disability care	111	989	1100 (4.1)
Mental health care	13	991	1004 (3.7)
General care and nursing (not ID-specific)	110	849	959 (3.5)
Physical disability care	35	768	803 (3.0)
	21,665	*5342*[a]	*27,007 (100)*

ID, intellectual disabilities.

[a] Distribution of 5342 primary diagnoses other than ID: psychiatric 2238 (41.9%), sensory disability 1061 (19.9%), physical disability 945 (17.7%), somatic disease 783 (14.7%), psychogeriatric 315 (5.9%).

Table 4
Distribution of people with Mild-ID support needs.

Eligibility	Total, n (%)
Single service use:	
Sheltered workplace only	49,861 (72.2)
Disability benefit only	11,631 (16.8)
Chronic care only	3033 (4.4)
Combined use of services:	
Sheltered workplace and disability benefits	2554 (3.7)
Sheltered workplace and chronic care	238 (0.3)
Disability benefit and chronic care	1643 (2.4)
User of all three services	118 (0.2)
	69,078 (100)

Distribution presented as observed after combining data from CCA (Chronic Care Act) and social benefits databases.
ID, intellectual disabilities.

level because of the difficulty in identifying individuals with ID consistently across different data sets.[20–22] The current method for data linkage and identification of ID groups can contribute to better retrieval of relevant information on the ID population. For example, this linkage method can provide a more accurate insight into the population size requiring ID-related public health services, keep track of developments over time and help to monitor effectiveness of interventions targeted at the ID population level. For specific future research investigations, the current method generates the largest possible cohort of Dutch individuals with ID.[23]

This study estimated a 1.45% ID prevalence in the Dutch adult population, which is in line with the literature. Previous studies mostly based their ID prevalence estimates on ID diagnoses, which suggests that the true ID prevalence might be closer to the upper limit (or even higher) rather than the lower limit of the generally reported prevalence estimates of between 1% and 1.5%.[14,24,25] In particular, this study found that in the younger age groups, ID prevalence rates were above 2.0% based on care use alone and above 3.5% when the Mild-ID group was included. In terms of the Mild-ID group, this study included individuals whose diagnosis is likely to be based on functional limitations or IQ scores only, rather than the conventional case definitions for ID, which also include limitations in adaptive functioning.[16,17] If the results were limited to individuals with ID receiving residential or non-residential care, the overall ID prevalence estimate amounts to 0.93%, which is slightly below the generally reported prevalence.[14]

While this investigation has successfully identified 67,078 individuals with a Mild-ID, other studies have previously suggested that, based on the normal distribution of IQ scores within a population, the total group of people with a Mild-ID in The Netherlands might be as large as 1.1 million people.[26] It is important to note that the underidentification of people with mild ID is a common problem of studies using administrative data.[11,22,24] Individuals with mild ID may not all rely on national systems for daily care and support, in contrast to those with more severe ID. Instead, supportive care to people with mild ID is more frequently organised by local and municipal programmes, of which the information is not included in national databases. In this study, we identified this group of individuals primarily through work- and income-related support systems; therefore, it could be seen as a sample of people with a mild ID and (some) capacity to do labour, albeit in a protective setting.

A limitation of using this type of administrative data is the absence of information about actual ID diagnoses. Although having an ID diagnosis is required to gain access to any of the services of which the data were used in this study, information about the diagnosis itself was not registered in any of the databases. For the administrative functioning of these databases, or the supportive systems themselves, detailed information on diagnoses is also not required; however, individuals with ID who are without a formally established diagnosis would not only be missing in the databases but might also lack access to services they require. Furthermore, the information available in this study showed which services individuals with ID were registered for, but not if these services were all used to the full extent as indicated. Information about diagnostic subgroups, for example, in DSM-V or ICD-10 classifications, would allow to further characterise subgroups, to identify any underrepresented groups and allow for international comparisons in a standardised manner.

A major strength of this study is the successful unique identification of people with ID at the national level. Although data were anonymised, the use of the unique identifier (RIN number) enabled the study to count users of multiple services as unique individuals across data sets. Moreover, the RIN number allows future linkage to a range of other data sets on health and well-being that do not necessarily require information about ID to be informative about individuals with ID. Examples can be found in our work on mortality, cancer and diabetes, which all rely on the methodology described in this study.[23,27,28] Both the national coverage and the RIN number therefore allow structural monitoring of trends in health and health care among people with ID and potentially also other vulnerable groups within Dutch supportive systems.[29]

Conclusions

This study successfully identified individuals with ID among users of national supportive services. The estimated ID prevalence in this study aligns with international estimates and shows a decline with increasing age. Databases from national supportive services therefore appear to be a useful resource to identify individuals with ID at the population level. As the current method is reproducible and relies on standard collected data, it provides opportunities for consistent monitoring of the Dutch population with ID in administrative data.

Author statements

Ethical approval

The study protocol for this exploratory study was reviewed by the Radboud University Medical Center institutional Ethics Committee who passed a positive judgment and waived the need for formal ethical assessment (2017–3921).

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Competing interests

Authors declare to have no conflicts of interest.

Data availability

Aggregated data from the databases used in this study are publicly available on a dedicated website of Statistics Netherlands (http://statline.cbs.nl). The non-public microdata used to link databases are, under certain conditions, accessible for statistical and scientific research (fees apply). Procedures can be found at www.cbs.nl, for further information: microdata@cbs.nl.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.puhe.2021.04.002.

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Short Communication

Paid sick leave benefits among essential frontline workers serving people experiencing homelessness in Canada during the COVID-19 pandemic

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ABSTRACT

Objectives: This study examined the prevalence and factors associated with paid sick leave benefits among direct service providers who work with people experiencing homelessness.
Study design: Cross-sectional study using an online survey disseminated during the second wave of the COVID-19 pandemic in Canada.
Methods: Survey data from 572 direct service providers working in the homeless, supportive housing, and harm reduction service sectors were analyzed for this study. Univariate and multivariate logistic regression models were used to examine predictors of paid sick leave benefits.
Results: One hundred one (17.7%) participants did not have any paid sick leave benefits. In the univariate models, paid sick leave was associated with older age, greater family income, full-time work, specific employment settings (supportive housing and not emergency shelters or harm reduction programs), having a regular medical doctor, and fewer occupational impacts of the COVID-19 pandemic. Older age, full-time work, and non-receipt of emergency financial benefits remained statistically significant predictors in the multivariate model.
Conclusions: Although the majority of service providers working with people experiencing homelessness have some amount of paid sick leave benefits, there is a precariously employed subset of individuals who are younger and working part-time in the sector. Temporary expansion of paid sick leave and removal of waiting periods for new employees to qualify for benefits are recommended.

Introduction

Paid sick leave has been identified as a key public health strategy for reducing transmission of COVID-19.[1] However, in Canada, over 40% of adults do not receive paid sick leave benefits from their employers.[2] In the context of the COVID-19 pandemic, this can force people to make difficult decisions between protecting public health or financially supporting themselves and their families.[3] Essential frontline workers without paid sick leave face the heightened precarity of being designated as needed by society (or, the laws enacted during the pandemic) and, thus, are expected to do their jobs without a protection that could keep them healthy and safe.[4] As some essential frontline workers serve vulnerable populations, the consequences of the lack of paid sick leave benefits can be deleterious not only to workers but also those receiving care and support.

People experiencing homelessness are one population that is highly vulnerable to COVID-19. A recent study found that people with recent histories of homelessness were over 3 times more likely to have a positive test result for COVID-19, 20 times more likely to be admitted to hospital for the virus, 10 times more likely to require intensive care for the virus, and 5 times more likely to die within the first 21 days of a positive test than people with stable

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housing.[5] One contributing factor to their heightened risk is the congregate nature of emergency shelters, which can be crowded and make other public health measures, such as physical distancing, more challenging.[6] Research has shown that these communal living settings are susceptible to outbreaks, which can cause rapid transmission of the virus to large proportions of service users and providers.[7,8] Yet, despite paid sick leave being an important public health intervention during the pandemic, there is a dearth of evidence on these benefits within the workforce that serves the homeless population. Understanding the prevalence and factors associated with paid sick leave benefits among this workforce could help to reduce COVID-19 transmission and spreading within the medically vulnerable homeless population. Accordingly, this study examined the prevalence and factors associated with paid sick leave among direct service providers who work with people experiencing homelessness in Canada.

Methods

As part of an ongoing study of the mental health and well-being of direct service providers working with people experiencing homelessness in Canada, an online survey was developed and disseminated widely across the country via emails to sector-specific networks and organizations, newsletters, listservs, social media, and word of mouth. The survey collected data on background and occupational information, including paid sick leave benefits; health and well-being; and perceived impacts of the COVID-19 pandemic. Service providers were eligible to participate in the study if they [a] were 18 years of age or older; [b] worked in Canada; [c] provided direct services to people experiencing homelessness; and [d] worked in homeless (including specialized health services for people experiencing homelessness), supportive housing, or harm reduction services. The study was approved by the institutional review board of the lead author's affiliation.

A total of 579 participants completed the online survey between November 12 and December 22, 2020—a six-week period during the second wave of the COVID-19 pandemic in Canada that had a daily average of 5949.07 cases and 91.46 deaths.[9] Seven (1.2%) participants were removed from analysis due to missing data on sick leave benefits. Univariate and multivariate logistic regression models were used to examine the predictors of paid sick leave benefits among direct service providers working with people experiencing homelessness. All variables that were statistically significant in the univariate models were entered into the multivariate model.

Results

The sample was similar to the homeless service sector workforce in gender (female: 79.9%; male: 16.6%; transgender/nonbinary: 2.8%), age (M = 39.05 years, SD = 12.60), ethnicity (white: 80.4; non-white: 18.4%).[10] Forty-four participants (7.7%) identified

Table 1
Predictors of paid sick leave benefits among 572 direct service providers working with people experiencing homelessness.

Predictor	Univariate models			Multivariate model		
	UOR	95% CI	*p*	AOR	95% CI	*p*
Gender						
Male (reference)	1.00					
Female/transgender/non-binary	0.69	0.37–1.30	0.25			
Age	1.04	1.02–1.06	**<.001**	1.04	1.02–1.07	**.001**
Ethnicity						
White (reference)	1.00					
Non-white/mixed race	1.06	0.61–1.86	0.83			
Education						
High school or less (reference)	1.00			1.00		
College diploma	2.34	1.07–5.13	**.03**	2.56	0.90–7.25	0.08
Bachelor's degree	1.74	0.83–3.62	0.14	2.34	0.86–6.37	0.10
Graduate degree	2.24	0.87–5.79	0.10	2.42	0.69–8.44	0.17
Family income						
$39,999 or less (reference)	1.00			1.00		
$49,000–59,999	5.14	2.66–9.95	**<.001**	2.11	0.91–4.92	0.08
$60,000–79,999	2.55	1.23–5.16	**<.01**	0.94	0.39–2.28	0.90
$80,000–99,999	2.69	1.31–5.52	**<.01**	1.31	0.52–3.28	0.57
$100,000 or more	3.24	1.73–6.04	**<.001**	1.35	0.58–3.14	0.49
Work amount						
Part-time (reference)	1.00			1.00		
Full-time	13.41	8.15–22.06	**<.001**	15.54	8.52–28.34	**<.001**
Lived experience of homelessness	1.10	0.61–2.01	0.75			
Lived experience of behavioral health problems	1.19	0.77–1.83	0.43			
Work setting						
Emergency shelter	0.57	0.36–0.89	**.02**	0.64	0.33–1.26	0.20
Supportive housing	2.03	1.18–3.51	**.01**	1.38	0.64–2.99	0.41
Community-based health	1.69	0.89–3.22	0.11			
Harm reduction	0.40	0.21–0.75	**<.01**	0.73	0.28–1.88	0.51
Have a regular medical doctor	1.75	1.02–3.00	**.04**	1.32	0.65–2.68	0.44
Unmet behavioral health need[a]	0.94	0.53–1.65	0.83			
COVID-19 impacts						
Worsened mental health	1.07	0.63–1.82	0.80			
Increased stress	1.78	0.88–3.58	0.11			
Increased alcohol use	0.83	0.51–1.33	0.43			
Increased cannabis use	0.83	0.50–1.37	0.46			
Decreased work hours	0.29	0.18–0.48	**<.001**	0.57	0.29–1.13	0.11
Accessed financial benefits[b]	0.26	0.16–0.44	**<.001**	0.35	0.17–0.72	**<.01**

UOR = unadjusted odds ratio; AOR = adjusted odds ratio; CI = confidence interval.

[a] Past 12 months.

[b] A national, emergency financial benefit for workers impacted by the pandemic.

as being Indigenous. Lived experience of behavioral health problems (57.0%) and homelessness (15.9%) were also common among direct service providers. All provinces and territories except the Yukon were represented in the sample; 9.1% were from the Atlantic region, 57.9% were from Central Canada, 15.1% were from the Prairie Provinces, 16.6% were from the West Coast, and 1.2% were from the Northern Territories. Most participants (70.1%) worked in urban settings, whereas 29.4% provided services to at least one rural, remote, or small community.

Of the 572 participants, 101 (17.7%) had no paid sick leave benefits, 121 (21.2%) had one week or less, and 350 (61.2%) had more than one week. The univariate and multivariate models predicting paid sick leave benefits (any amount) are shown in Table 1. In the univariate models, paid sick leave benefits were positively associated with older age, greater family income, full-time work, employment in supportive housing, and having a regular medical doctor. By contrast, service providers without paid sick leave benefits were more likely to be working in emergency shelters and harm reduction programs, have had their work hours decreased during the COVID-19 pandemic, and have accessed pandemic-specific emergency financial benefits (i.e. Canada Emergency Response Benefit). In the multivariate model, three factors remained significant predictors of paid sick leave benefits: older age, full-time employment, and non-receipt of emergency financial benefits.

Discussion

The results highlight that most service providers working with people experiencing homelessness have some amount of paid sick leave benefits. However, there is a precariously employed subset of service providers who are younger and working part-time in the sector. This group was also more likely to have accessed governmental emergency financial benefits, suggesting that the pandemic has had a greater financial toll on those without paid sick leave. As emergency shelters and harm reduction programs were negatively associated with paid sick leave benefits in the univariate models, but not the multivariate one, this is likely due to these settings having a higher proportion of part-time workers (emergency shelters: 22.6% part-time workers; harm reduction programs: 35.3% part-time workers; full sample: 19.0% part-time workers). Given that these services provide support to groups that are medically vulnerable, policies are needed to temporarily expand paid sick leave benefits to service providers working in these settings for the duration of the pandemic. Furthermore, as emergency shelters are vulnerable to COVID-19 outbreaks, preventive measures, such as paid sick leave, are critical to protecting service users and providers. Similarly, removal of waiting periods to qualify for paid sick leave benefits for new employees is strongly recommended. Those without paid sick leave in this study were also more likely to draw on these financial benefits. Given this, providing paid sick leave to the precariously employed subset of service providers without benefits could potentially offset the costs of governmental emergency financial benefits.

There are several notable study limitations. First, few participants were from Quebec, so the findings may be less generalizable to service providers working in that province. Second, the sample was more educated than the homeless service sector workforce.[10] Third, as the online survey was disseminated via homeless service networks and organizations, part-time employees without paid sick leave benefits, especially those who had their work hours reduced, may have been less able to participate in the study. The latter two limitations may have contributed to a higher rate of paid sick leave benefits in the sample. Nevertheless, the findings highlight a critical gap in paid sick leave benefits within an essential frontline workforce that needs to be addressed to reduce risks of COVID-19 transmission and spreading within the homeless population.

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Competing interests

None declared.

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Short Communication

Resistance to COVID-19 vaccination has increased in Ireland and the United Kingdom during the pandemic

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ABSTRACT

Objectives: Hesitance and resistance to COVID-19 vaccination poses a serious challenge to achieving adequate vaccine uptake in the general population. Cross-sectional data from the early months of the pandemic indicates that approximately one-third of adults in multiple nations are hesitant or resistant to a vaccine for COVID-19. Using longitudinal data, we tracked changes in attitudes to COVID-19 vaccination during the pandemic.
Study design: This is a quantitative, longitudinal design.
Method: Nationally representative samples of the adult general population of the Republic of Ireland (N = 1041) and the United Kingdom (N = 2025) were assessed for their attitudes towards COVID-19 vaccination at three points from March to August 2020.
Results: Statistically significant increases in resistance to COVID-19 vaccination were observed in Irish (from 9.5% to 18.1%) and British (from 6.2% to 10%) adults.
Conclusion: Resistance to vaccination has significantly increased in two European nations as the pandemic has progressed. Growing resistance to COVID-19 vaccination will pose a challenge to public health officials responsible for ensuring sufficient vaccine coverage.

With an excess of 65 million confirmed cases across 190 countries globally and more than 1.5 million deaths, COVID-19 remains one of the most important global health challenges of our lifetime.[1] Recent announcements about the development of multiple, safe and effective vaccines for COVID-19 offer a promising solution to this public health crisis. However, in addition to production and distribution challenges, a key challenge remains convincing a critical mass of the population to accept a COVID-19 vaccine.

In June 2020, 72% of people from 19 countries indicated that they would accept a COVID-19 vaccine, with rates as high as 89% in China and as low as 55% in Russia.[2] An ongoing Gallup Poll of the United States general population found that acceptance of a COVID-19 vaccine fell from 66% in July to 50% in September before increasing to 58% by the end of October.[3] Although more longitudinal research is clearly required, early evidence suggests that people's attitudes towards a COVID-19 vaccine are fluctuating over time.

In March 2020, our research group initiated a longitudinal project that tracked nationally representative samples of adults from the Republic of Ireland and the United Kingdom (UK). Asked if they would accept a hypothetical vaccine for COVID-19, approximately two-thirds of Irish (65%) and British (69%) adults responded ‘yes’, about one-quarter responded ‘maybe’ (26% in Ireland and 25% in the UK) and the remainder responded ‘no’ (9% in Ireland and 6% in the UK).[4] Here, we report changes in COVID-19 vaccine acceptance (‘yes’), hesitance (‘maybe’) and resistance (‘no’) in these samples during the first four months of the pandemic.

Identical sampling methods were used to collect data in Ireland and the UK. Participants were recruited from existing survey panels

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Table 1
Proportion of Irish (n = 1030) and UK (n = 2020) samples accepting, hesitant to and resistant to a COVID-19 vaccine.

Country	Wave 1	Wave 2	Wave 3	Pairwise comparisons (Wald chi-squared)		
	% (95% CI)	% (95% CI)	% (95% CI)	Wave 1 vs wave 2	Wave 1 vs wave 3	Wave 2 vs wave 3
Ireland						
Acceptance	64.9 (61.9, 67.8)	64.3 (60.5, 68.1)	55.4 (51.6, 59.3)	Chi-squared = 0.09	Chi-squared = 24.67***	Chi-squared = 22.05***
Hesitance	25.6 (22.9, 28.3)	23.0 (19.5, 26.4)	26.6 (23.0, 30.2)	Chi-squared = 2.02	Chi-squared = 0.23	Chi-squared = 3.08
Resistance	9.5 (7.7, 11.3)	12.9 (10.1, 15.6)	18.1 (15.0, 21.2)	Chi-squared = 6.27*	Chi-squared = 30.24***	Chi-squared = 11.03***
UK						
Acceptance	68.8 (66.8, 70.9)	66.6 (64.3, 69.0)	71.8 (69.4, 74.3)	Chi-squared = 3.12	Chi-squared = 5.53*	Chi-squared = 18.97***
Hesitance	24.9 (23.1, 26.8)	23.8 (21.7, 26.0)	18.1 (16.0, 20.2)	Chi-squared = 0.88	Chi-squared = 30.29***	Chi-squared = 23.54***
Resistance	6.2 (5.2, 7.3)	9.4 (7.9, 10.9)	10.0 (8.4, 11.7)	Chi-squared = 17.68***	Chi-squared = 20.57***	Chi-squared = 0.51

95% CI = 95% confidence interval; UK, United Kingdom.
All degrees of freedom = 1
* $P < .05$.
** $P < .01$.
*** $P < .001$.

by Qualtrics using quota sampling to ensure the samples were representative of their populations by sex, age, regional and income distributions.[4,5] In Ireland, wave 1 ($N = 1041$) longitudinal data were collected from March 30 to April 5, wave 2 ($n = 506$, recontact rate = 49%) data were collected from April 30 to May 19 and wave 3 ($n = 534$, recontact rate = 51%) data were collected from July 16 to August 8. In the UK, wave 1 ($N = 2025$) longitudinal data were collected from March 23 to 28, wave 2 ($n = 1,406$, recontact rate = 69%) data were collected from April 22 to May 1 and wave 3 ($n = 1,166$, recontact rate = 58%) data were collected from July 9 to August 9.

At each assessment, participants were asked ‘If a new vaccine were to be developed that could prevent COVID-19, would you accept it for yourself?’ Response options were ‘yes’, ‘maybe’ and ‘no’. Changes in COVID-19 vaccine acceptance (‘yes’), hesitance (‘maybe) and resistance (‘no’) were assessed using a structural equation modelling framework modelling, and missing data were handled using robust maximum likelihood estimation.[6,7] A ‘null’ model was initially specified wherein the proportions (e.g., in vaccine resistance) were constrained to be equal over time. Next, an ‘alternative’ model was specified wherein the proportions were freely estimated over time. These models differ by one degree of freedom, so improvement in model fit can be tested using a log-likelihood ratio test that is distributed as a chi-squared distribution. Finally, pairwise comparisons were tested using a Wald chi-squared test. All analyses were performed in Mplus, version 8.2.[8]

There were statistically significant changes over time in attitudes towards COVID-19 vaccination in the Irish (chi-squared (2, 1030) = 33.37, $P < .001$) and British (chi-squared (2, 2020) = 19.22, $P < .001$) samples. Table 1 reports all figures including the pairwise comparisons. In Ireland, between March and August 2020, there was a significant decrease in vaccine acceptance (from 64.9% to 55.4%), no change in vaccine hesitance (25.6–26.6%) and a significant increase in vaccine resistance (9.5–18.1%). During the same period in the UK, there was a significant increase in vaccine acceptance (68.8–71.8%), a significant decrease in vaccine hesitance (24.9–18.1%) and a significant increase in vaccine resistance (6.2–10.0%).

Substantial changes in attitudes towards a vaccine for COVID-19 were evident in Irish and British adults during the first four months of the pandemic. Resistance to COVID-19 vaccination rose in Ireland by 91% and in the UK by 61%. It was notable that the changes in attitudes towards COVID-19 vaccination changed in distinct ways in the two countries. The rise in vaccine resistance in Ireland was associated with a concomitant fall in vaccine acceptance, whereas in the UK, the rise in vaccine acceptance and resistance was associated with a parallel fall in vaccine hesitance.

The levels of vaccine acceptance identified in Ireland and the UK during the earliest phase of the pandemic were similar to the average level of vaccine acceptance across 19 nations reported by Lazarus et al.[2] in June 2020. However, the increased rate of vaccine resistance observed in both countries by August 2020 mirrors similar trends observed in the United States at a similar period of time.[3] It appears that the rise in vaccine resistance coincided with the loosening of public health restrictions in these countries during the summer months as the spread of the virus began to come under control and normal daily functioning for many people resumed. The emergence of the ‘second wave’ of COVID-19 appears to have led to an increase in vaccine acceptance in the United States,[3] and it will be important to understand if similar trends are observed in other nations.

Moreover, the emergence and public distribution of actual vaccines for COVID-19 may shift public attitudes further, which presents an opportunity to assess how COVID-19 vaccine acceptance, hesitance and resistance change in the context of an available vaccine; what predicts change in attitudes towards vaccination and how well attitudes towards a hypothetical vaccine predict actual vaccine uptake. Timely, evidence-based answers to these questions will provide valuable public health information to health officials and policymakers to better inform social behaviour and communications strategies to increase the uptake of a COVID-19 vaccine.

Author statements

Ethical approval

Collection of the data presented in this study was approved by the Research Ethics Committees at The University of Sheffield and Ulster University.

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Competing interests

All authors have no conflict of interest to declare.

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