Adverse Mortality Trends Have Increased the Number Bereaved in the United States

Leah R. Abrams[1,*], Sha Jiang[2,*,†], Diego Alburez-Gutierrez[2], Enrique Acosta[3], and Emilio Zagheni[2]

[1]Tufts University Department of Community Health, Medford, MA, USA
[2]Max Planck Institute for Demographic Research, Rostock, Germany
[3]Centre d'Estudis Demogràfics, Barcelona, Spain
[*]Co-first authors
[†]Corresponding author: jiang@demogr.mpg.de

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Abstract

U.S. life expectancy plateaued since 2010 and declined in 2014-2017 and 2020–2021, increasing the gap between the United States and its peers. This analysis calculated U.S. trends in the number of individuals who lost a close kin, such as a mother, father, at least one sibling, or at least one child, and how the number of bereaved people changes under counterfactual mortality conditions.

We found that adverse mortality in 2010–2019 and spikes in mortality in 2020–2021 resulted in a substantial increase in the number of individuals bereaved annually. Increases over time were especially large for individuals bereaved by sibling loss.

If the United States had maintained its rates of mortality improvement from 2000–2009, about 2.7 million fewer people would have experienced at least one close-kin loss during 2010–2019. By 2021, the excess number of individuals who had ever been bereaved since 2010 had grown to 5.6 million. Considering single years, in 2019 alone 8.6% of all bereaved individuals—around 0.9 million people—would not have lost any close relative had the United States maintained its previous pace of mortality improvement. In 2021, the corresponding annual excess rose to 25%, or about 3.2 million individuals. Excess bereavement is similar when applying the concurrent rates of mortality improvement of countries like Japan and Switzerland. Decomposition analyses revealed that excess bereavement in the United States can be explained primarily by adverse mortality trends, rather than by differences in size or composition of the population at risk.

Considering population health trends from the perspective of the number bereaved shifts the focus from individual deaths to their rippling effects on surviving family, at a population scale.

Significance Statement

Since 2010, the United States has been trailing peer countries in life expectancy improvements. We show that adverse U.S. mortality trends have resulted in a substantial increase in the number of people annually losing their mother, father, a sibling, or a child. Increases over time have been especially large for sibling loss. If U.S. rates of mortality improvement from 2000–2009 had continued in 2010–2019, almost 9% of bereaved individuals in 2019—0.9 million people—would not have experienced bereavement. In 2021, these figures rose to 25%, 3.2 million. Calculating trends in bereavement shifts the focus of population health research from individual deaths to their rippling effects on surviving family members, highlighting the importance of grief support and returning mortality declines.

Main Text

Introduction

After decades of steady improvement, life expectancy in the United States plateaued since 2010 [1], markedly decreased in 2014–2017 [2] and 2020–2021 [3, 4], and by 2023, had not yet fully recovered [5]. Before the COVID-19 pandemic, the lack of improvement in U.S. life expectancy could be explained by increases in deaths by suicide, homicide, and drug overdoses, along with stagnation in declining cardiovascular disease mortality [6, 7]. These mortality trends and the COVID-19 pandemic were not unique to the United States, but their severity and their impact on life expectancy were outsized, resulting in a growing gap between U.S. life expectancy and that of peer nations [1, 3].

While U.S. mortality trends and international life expectancy comparisons have been established in the scientific literature, including a National Academies of Sciences report on the alarming adverse trends in mortality in midlife [8], life expectancy remains an elusive measure to much of the public and policymakers. For example, life expectancy holding steady around age 78.8 in 2010–2019 does not immediately convey that more Americans were dying than prior trends would suggest, and at younger ages. In addition to the fact that life expectancy is often based on a synthetic cohort and thus does not provide the real expected life span for anyone, the implicit age standardization in life expectancy calculations obscure the role of changes in age-specific mortality [9]. Beyond these technical problems for interpretation, measuring life expectancy does not capture the impact of mortality trends on the individuals, families, and communities surviving loss.

Population trends in bereavement are another way of measuring the impact of mortality trends that may be easier to translate to the public and policymakers, with clearer implications. Prior research has examined population levels of bereavement due to specific causes, such as COVID-19 [10, 11, 12] or drug overdose deaths [13], and specific types of bereavement, such as mothers losing children [14] and children losing co-residential parents [15]. From this research, we have learned, for example, that each COVID-19 death resulted in approximately nine bereaved persons in the United States who lost a grandparent, parent, sibling, spouse, or child [12]. Research on population bereavement has also documented disparities between groups given persistent social inequalities in mortality and how social forces contribute to bereavement experiences [16, 17]. A recent study in this area revealed that Black and Native American older adults report more premature family loss and a greater cumulative loss of family over their lifetime [18]. Taken together, the existing literature in this area has provided snapshots of the impact of mortality on population bereavement, often segmented by decedent's relationship and death timing [16], but has not provided insights into the total impact of long-term adverse mortality trends from multiple different causes of death on the number of people experiencing bereavement over time.

Doing so is important to communicating the impact of population health trends on residents of the United States. Bereavement, and especially untimely or unexpected loss, is associated with poor mental health [19, 20], physical health [20], behavioral outcomes [19, 21, 22], economic outcomes [23], and sometimes even subsequent mortality risk [20, 24, 25, 18]. Beyond individuals, losses impact family dynamics and the functioning of family systems [26, 27, 28] and can disrupt entire communities [29, 30, 31]. There is some evidence that untimely loss, like early loss of a parent, and unexpected loss, like deaths due to accidents, suicides, or natural disasters, can have especially long impacts on those grieving [32, 33]. Premature losses may have distinct impacts on the bereaved, including higher subsequent mortality risk, because they are more likely to be perceived as preventable, stigmatizing, and unresolved [18].

Despite strong evidence of the detrimental impact of bereavement on multiple dimensions of well-being, there remains little institutional support for those grieving in the United States, as the Family and Medical Leave Act (FMLA) does not explicitly provide bereavement leave, and coverage provided by states or employers is often inadequate [34]. Even in the cases of expected loss, bereaved families often face challenges in accessing critical psychosocial support [35]. Calculating population trends in the number of bereaved individuals in the population brings attention to survivors and their needs for support as a public health priority while motivating researchers and policymakers to work towards improvements in population mortality.

Therefore, in this study, we use demographic kinship models to answer the following research questions: How has the number of individuals losing at least one close kin at critical ages changed over time in the United States? We consider close kin to be mothers, fathers, siblings, and children, examining losses early in life, across the life course, and late in life. We estimate the annual number of individuals experiencing kin lossand the cumulative number of ever-bereaved individuals, which aggregates across years and different types of kin [36]. Then we ask: How many of these bereaved individuals would not be bereaved if the United States had maintained earlier rates of mortality improvements or had the mortality improvements of peer nations? Such counterfactual analyses estimate "excess" or "preventable" bereavement. We model bereaved persons, rather than the number of losses, to move from the mortality orientation focused on individual deaths to a bereavement framework focused on the experiences of individuals in need of grief support.

Results

Trends in the number and proportion bereaved

The annual number of individuals who lost a mother, father, at least one sibling, or at least one child has been steadily rising in the United States from 1970–2021, with slopes increasing in 2010–2019 and more so in 2020–2021. Figure 1 shows these trends in the number of bereaved individuals and the proportion of the population bereaved from each type of close kin in critical age groups. In 1970–2009, when mortality in the United States was robustly improving, the annual number of people bereaved from each type of kin loss was already increasing, more so than the proportion of the population bereaved annually. The growth in the proportion bereaved was driven mainly by kinship structure: as survival improved and fertility remained relatively stable, more relatives lived to older ages, raising individuals' probability of experiencing at least one kin loss. By contrast, the growth in the absolute number bereaved also reflected population components: mortality decline also resulted in increased total population size and population aging, meaning more individuals were "available" to experience kin loss. As a result, the number bereaved began to rise earlier and more steeply than the proportion bereaved. In 2010–2019, when adverse mortality trends held back life expectancy improvements, both the number and the proportion bereaved annually increased in most cases. The annual numbers of individuals who lost a mother, father, or at least one sibling at any age (0–110) or lost at least one child in/after fertile ages (15–110), steadily rose, as did the annual proportions of the population with these types of loss. In some cases, such as the annual number losing at least one sibling at any age, the increasing slope is noticeably steeper in 2010–2019 (rising from 2805 to 3594 thousand, with an average annual increase of approximately 79 thousand) than in 1970–2009 (rising from 1917 to 2766 thousand, with an average annual increase of approximately 21 thousand). The annual *number* of individuals who lost a parent or sibling in the critical ages of 0–18—a rare event—was relatively stable, although the annual *proportion* having lost a mother or father in childhood was increasing slightly in 2010–2019. The annual number bereaved from child loss in any age and in old ages (65–110) also rose in 2010–2019, even though the proportion of the older population with child loss was falling, given the growing older population. In 2020–2021, when mortality rose sharply in the United States, the number bereaved and the proportion bereaved for every type of kin loss spiked, to varying degrees.

Figure 2 shows the age distributions of the annual number bereaved and the annual proportion of the population bereaved from each type of kin loss in 2009 (baseline), 2019 (after a decade of adverse mortality trends), and 2021 (in the height of COVID-19 related mortality). For all four of these types of close family loss, across most of the life course, the proportion and the number bereaved were highest in 2021, followed by 2019, and lowest in 2009. The differences are largest for sibling loss, which peaked around age 62 at 77 thousand in 2009, age 72 at 114 thousand in 2019, and age 66 at 147 thousand persons bereaved in 2021. Losing a father occurred at younger ages on average than losing a mother. For example, at age 30, 52 thousand people in the United States had lost a father in 2021, compared to 40 thousand in 2019, and 35 thousand in 2009. We do not observe similar numbers bereaved from mother loss until around age 42. While Figure 2 shows large differences in the proportion of the older population bereaved by child loss annually, differences over time in the number who experienced child loss were less substantial.

Characterizing excess bereavement

Our counterfactual analysis calculated the excess number of bereaved individuals in the United States each year by comparing baseline (observed) bereavement to the number that would have been bereaved if the United States had maintained its rate of mortality improvement (ROMI) from 2000–2009 in 2010–2021 or if the United States had the ROMIs of Italy, Japan, Switzerland, and the United Kingdom in 2010–2021 (Figure 3). Around 2.7 million individuals experienced loss of at least one close kin (mother, father, sibling, or child) in 2010–2019 that would have been prevented if the United States had followed historical mortality trends of 2000–2009 in 2010–2019. About 0.9 million individuals would not have experienced any close kin loss in 2019 alone had the United States maintained prior mortality improvements, representing 8.6% of that year's bereaved population. These calculations refer to the number of people spared the experience of ever losing close kin. Such estimates can be considered conservative in that they do not account for changes in the intensity of bereavement for those who lost multiple close kin.

Looking by kin type, excess bereavement in the United States compared to counterfactual scenarios was most notable for losing a sibling at all ages, followed by losing a father or mother. For example, 353 thousand fewer people in the United States would have experienced the loss of at least one sibling in 2019 if the United States had maintained its ROMI from 2000–2009 in 2010–2019 (translating to a roughly 10% reduction in individuals bereaved from sibling loss). In addition, around 65 thousand fewer children (aged 0–18) in the United States would have lost a father in 2019 had U.S. mortality been improving at the rate of Switzerland in 2010–2019, representing around 30% fewer bereaved children in 2019. To a lesser extent, there was also excess bereavement from child loss in the United States. These patterns of kin loss are likely connected, as adverse mortality trends in middle-aged men could contribute to people losing a sibling in middle age, losing a father during childhood, or losing a child at old ages.

During the COVID-19 pandemic in 2020–2021, excess bereaved persons in the United States increased dramatically, consistent with evidence that the United States experienced disproportionate increases in mortality during the pandemic compared to other high-income countries. If the U.S. mortality improvements from 2000–2009 had continued through 2021 (eliminating the mortality impacts of the COVID-19 pandemic), about 1.3 million fewer individuals would have lost at least one sibling at any age in 2021, which represents 28% of individuals bereaved by sibling loss that year. This number is similar to the number of excess bereaved persons under the perhaps more realistic counterfactual scenario applying the ROMIs of Japan and Switzerland from 2020–2021.

The differences are relatively smaller when considering the ROMIs of Italy and the United Kingdom during these pandemic years. Some scenarios show dramatic comparisons; for example, the United States would have had around 50% fewer individuals aged 0–18 experiencing father loss in 2021 if it had the ROMI of Japan or Switzerland. Considering individuals who lost at least one close kin (mother, father, sibling, or child) at any time between 2010 and 2021, there were a total of 5.6 million excess bereaved individuals compared to what would have occurred if U.S. mortality trends of 2000–2009 had continued. In 2021 alone, 3.2 million more individuals lost at least one close family member during that year than would have been bereaved under historical trends (25% more than observed in baseline 2021).

Decomposition analysis revealed that excess bereavement in the United States in 2015, 2019, and 2021 can be explained almost entirely by differences in bereavement probabilities with an average contribution of around 130% across different ages of bereaved persons and different kin types, fueled by adverse U.S. mortality trends, rather than differences in the size and distribution of the population at risk of becoming bereaved which has negative average contribution of around 30% (Figure 4). Similar decomposition patterns can be observed for individuals in specific age groups across years in Figure C3. Difference in bereavement probabilities contributed to excess bereavement across the life course in similar age distributions to those seen in the number bereaved in Figure 2A. Specifically, the bereavement component makes the largest contributions to excess mortality around midlife (ages 40–60) for losing a mother or father, while mortality at older ages (60–80 years old) appears more important for losing at least one sibling or child. The population component shows not the effect of population size and age structure itself, but the difference between U.S. size/structure and counterfactual size/structure, where the latter is calculated by applying counterfactual mortality while keeping all other demographic processes at the U.S. level. Adverse mortality trends in the United States reduce the population size (meaning fewer individuals available to experience kin loss), so despite population aging, the population component makes a small negative contribution to excess bereavement.

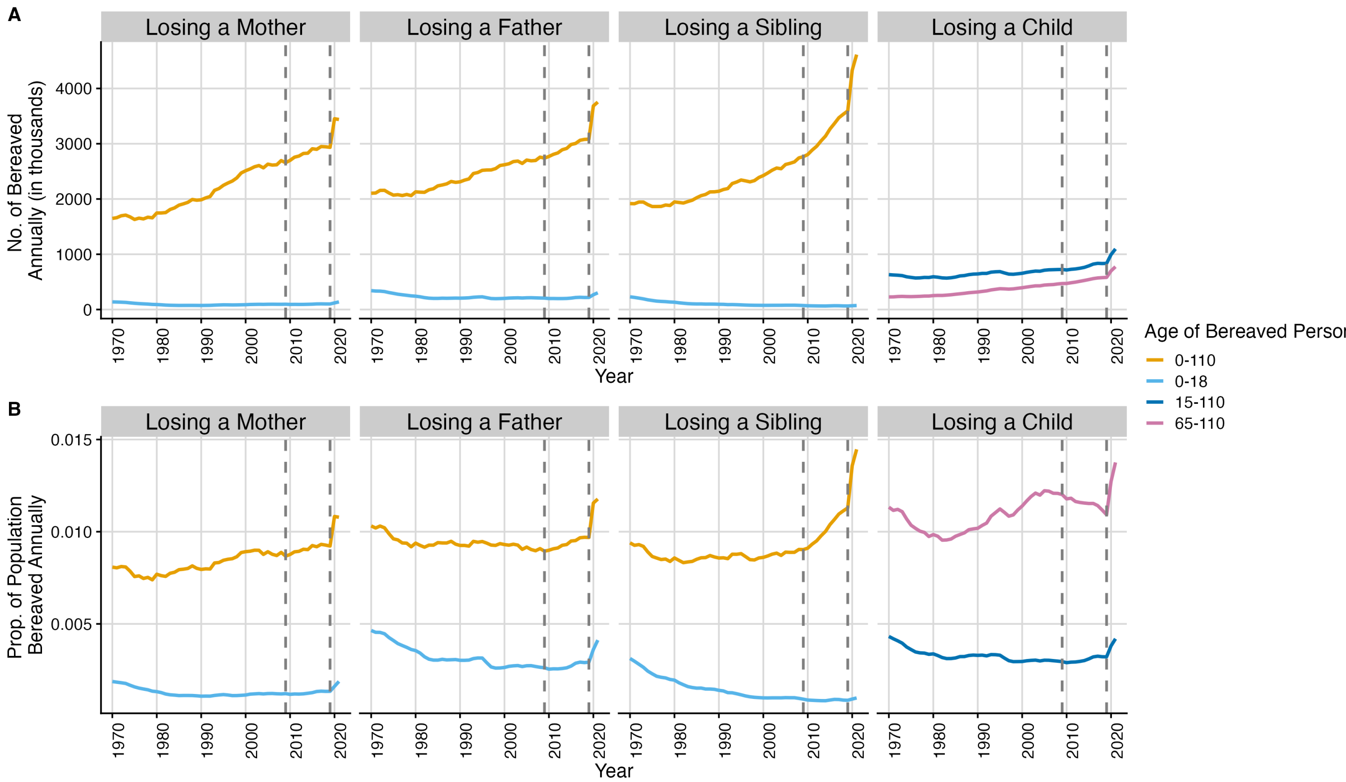

Figure 1: Number of individuals bereaved annually by close kin loss (in thousands, row A) and the proportion of population by age group experiencing close kin loss annually (row B) from 1970 to 2021. Columns show the corresponding results of losing different types of kin. Colors represent different age groups of bereaved persons. Dashed gray reference lines mark the years 2009 and 2019.

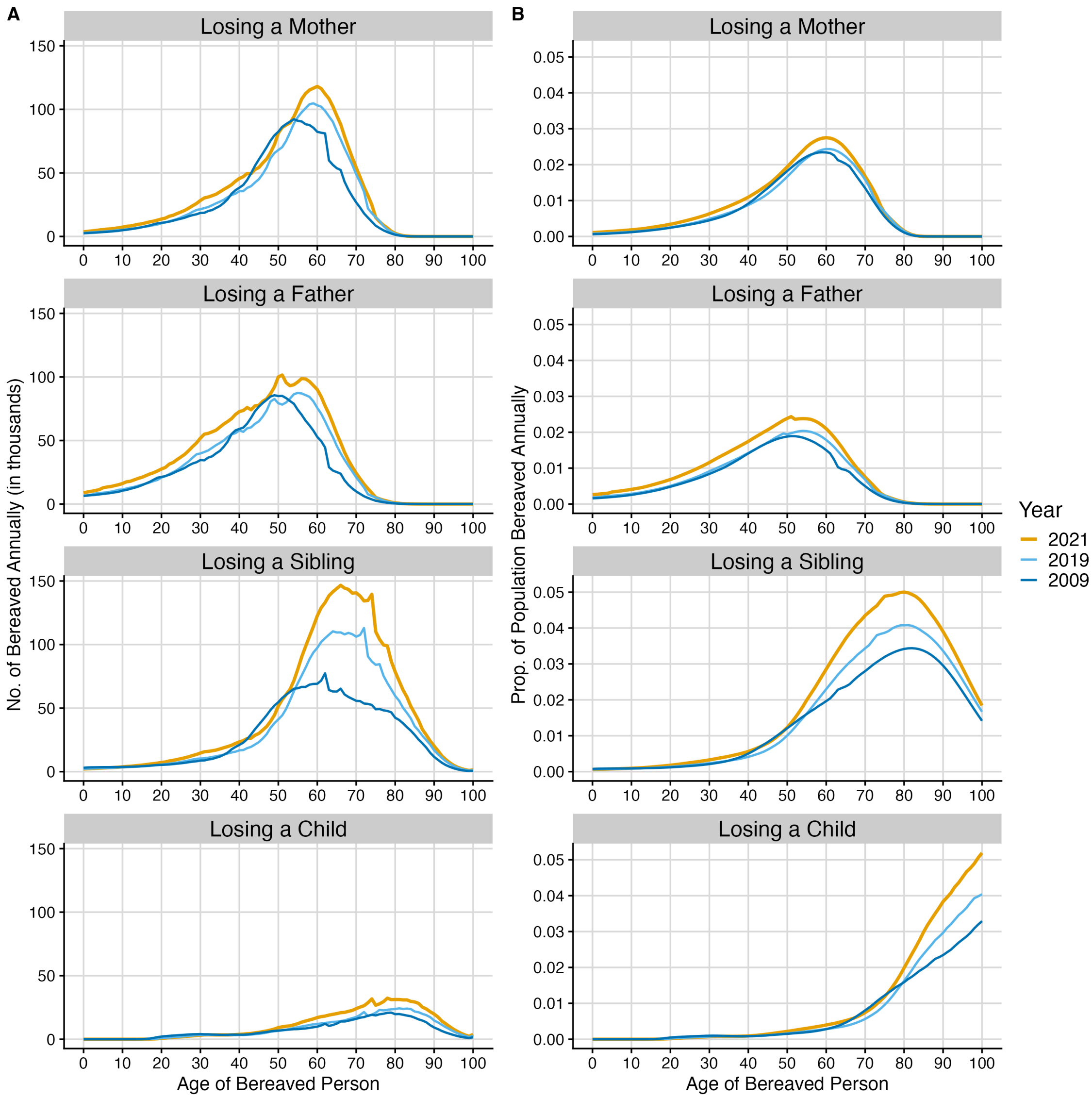

Figure 2: Number of individuals bereaved annually by close kin loss (in thousands, column A) and the proportion of population experiencing close kin loss annually (column B) across ages. Different rows show the probability of losing different types of kin, while line types show different years.

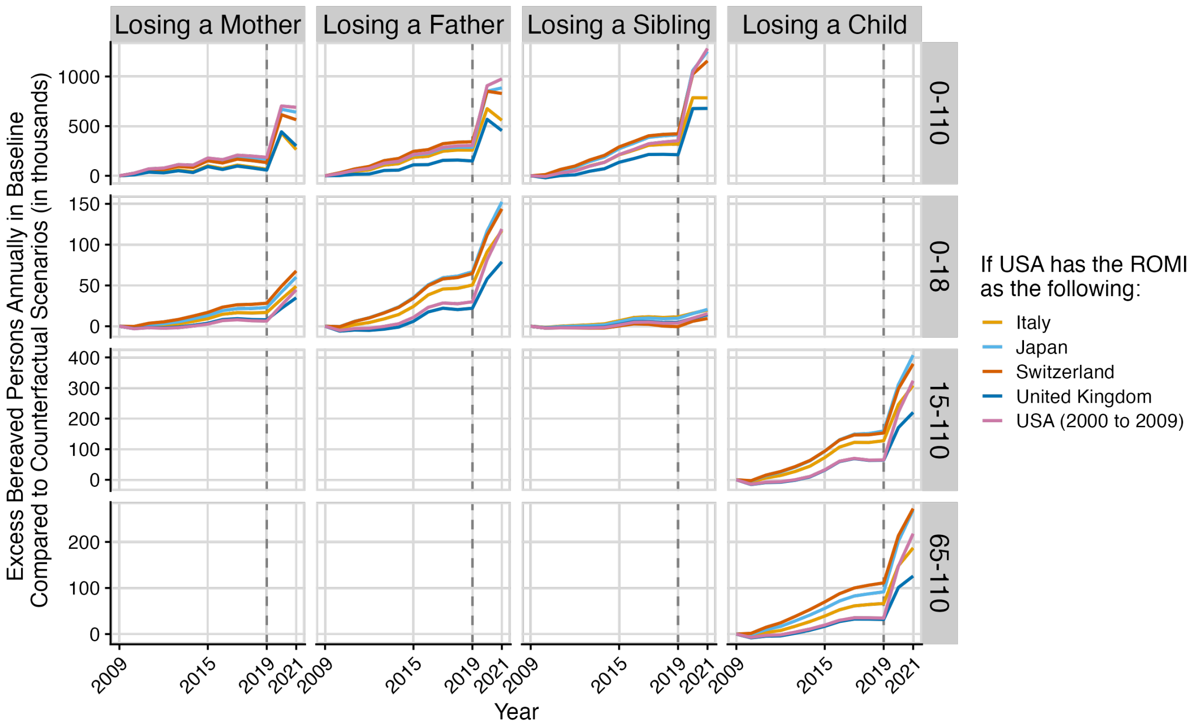

Figure 3: Difference in the number of individuals bereaved by close kin loss annually from 2009 to 2021, calculated as baseline minus each counterfactual scenario (in thousands). Here, counterfactual scenarios assume that the United States had the rate of mortality improvement (ROMI) of peer nations in 2010–2021, or it had maintained its ROMI from 2000–2009 in 2010–2021. Colors represent the difference between the baseline and different counterfactual scenarios. Results are shown by bereaved individuals of different age groups (by rows) and by the different types of kin they lose (by columns). Dashed gray line marks 2019.

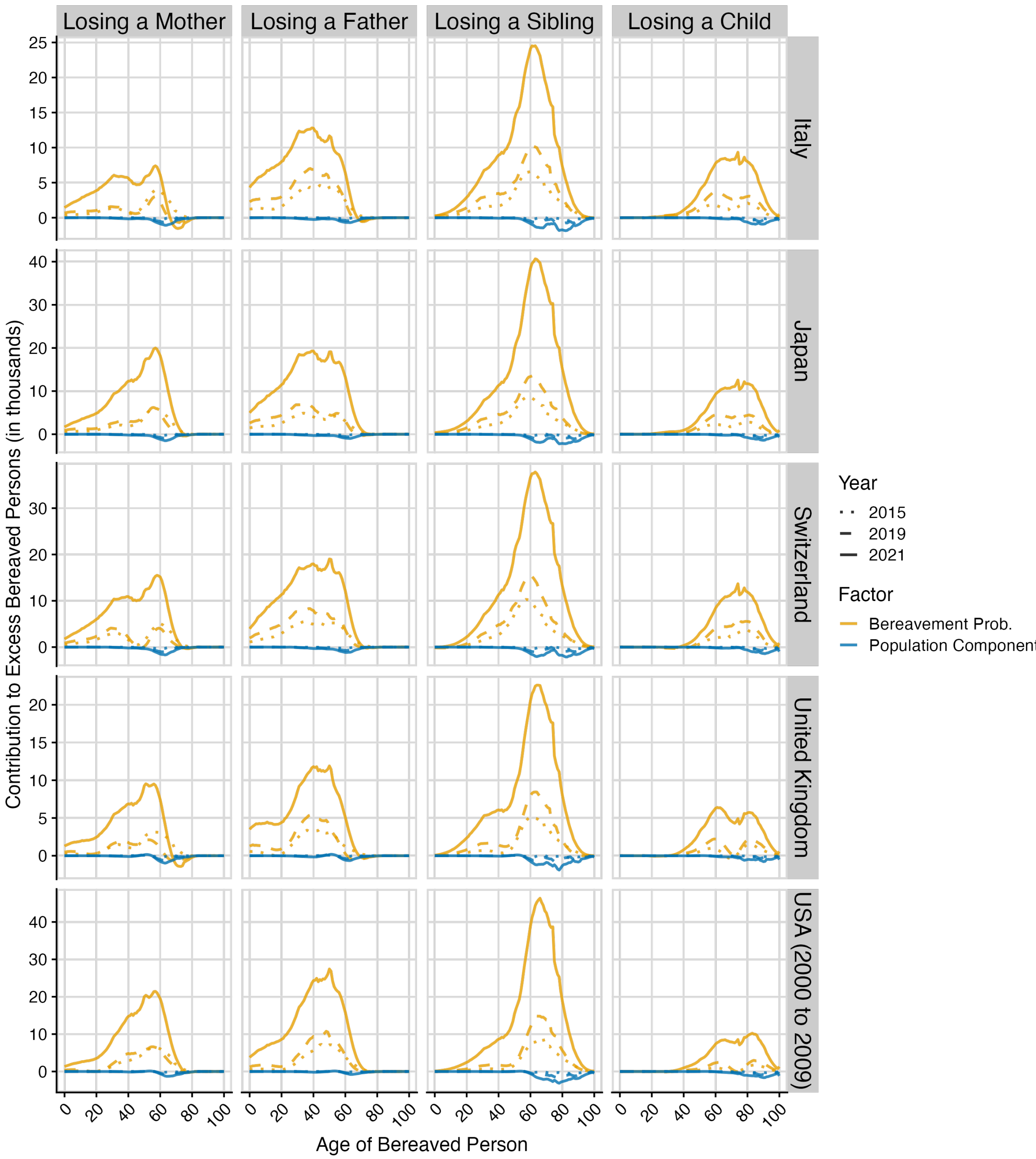

Figure 4: Contributions of bereavement probability and population component to the excess bereaved persons (in thousands) across ages of bereaved persons. Such excess bereaved persons are calculated as the difference in bereaved persons between baseline and counterfactual scenarios. Colors represent the contribution of different factors. Lines show different years. Results are shown by different counterfactual scenarios (by rows) and by the different types of kin they lose (by columns).

Discussion

Adverse mortality trends in the United States in 2010–2019 and spikes in U.S. mortality in 2020–2021 resulted in a substantial increase in the number of individuals bereaved annually due to losing a mother, father, at least one sibling, or at least one child. In 2019, nearly 9% of individuals who experienced at least one close-kin loss that year could be considered excess or preventable bereavement that would not have occurred if the United States had maintained mortality declines of 2000–2009. In 2021, 23% of individuals bereaved in that year would not be grieving the loss of any close family member if the United States had the mortality improvements of Japan or Switzerland in 2020–2021 (or around 25% if the United States had continued its previous pace of mortality improvement). Decomposition analysis confirmed that this excess bereavement in the United States since 2010 is explained by adverse mortality trends and not population size and structure differences.

Our findings are consistent with other studies that have shown that the gap between life expectancy in the United States and in other high-income countries has been growing since 2010, and especially during the COVID-19 pandemic [1, 3]. Aburto et al. [3] reported that men in Switzerland had the highest life expectancy at birth in 2019 of any of the 29 observed countries at 82.2 years, 5.5 years higher than the life expectancy of men in the United States that year. Our analysis shows how the mortality conditions that produced this large gap in life expectancy can be thought of in terms of the number of people subjected to bereavement and its accompanying grief — over one million more individuals in the United States bereaved due to loss of close kin in 2019 than would have occurred with Switzerland's rate of mortality improvement in 2010–2019. Woolf [1] reported that COVID-19 mortality resulted in the United States falling further in international life expectancy rankings, from 40th in 2019 to 46th in 2020. Our bereavement analysis shows there were 1.8 million more people bereaved by close kin loss in 2020 compared to 2019 and 22.9% preventable cases in 2020 that would not have occurred under the mortality trends of Japan.

Translating the U.S. mortality disadvantage into numbers and proportion bereaved offers an alternative, interpretable, and affecting metric for conceptualizing how adverse trends in population health impact families, communities, and the nation at large. In addition, doing so reveals new implications that might otherwise be missed. For example, between 2009 and 2021, there was a dramatic increase in the number of individuals in the United States and the proportion of the U.S. population who lost at least one sibling, especially between ages 50–80 years old. Sibling loss is sometimes considered an under-recognized, disenfranchised form of grief [37, 38], and much of the literature on this experience focuses earlier in the life course (for example, [39, 40, 41, 23]). Losing a sibling in midlife may have unique implications, such as increasing pressure on surviving siblings to provide care to aging parents, that warrant further investigation.

The psychology literature has thoroughly reported the emotional and mental health outcomes following bereavement events, and epidemiological studies have documented alarming health risks [16, 20]. Even so, there remains a dearth of institutional support for bereavement in the United States. The lack of federal policy granting paid or unpaid time away from work following a loss leaves grieving workers at the whim of their employer or manager if they need time off while mourning. If individuals are not diagnosed with prolonged grief disorder, they may have difficulty obtaining insurance coverage for mental health services such as grief counseling [42]. Research has shown that a large share of bereaved individuals who meet criteria for prolonged grief disorder, major depressive disorder, post-traumatic stress disorder, and other mental health conditions still do not have access to mental health services, suggesting additional barriers such as awareness of maladaptive symptoms, connecting with an appropriate health care provider, and stigma [43]. Our results suggest that such individual experiences and challenges have been increasing over time in the population and are happening in excess on the population scale compared to what might have been expected based on historical trends. In addition, our study brings attention the potential of an intergenerational spiral in which increasing premature death can impact the social, economic, mental, and physical well-being of survivors in the same or next generation.

These implications should be considered along with the limitations of this study. The analysis does not account for immigration or deaths that occur outside of the United States. Doing so would require information on immigration rates, countries of origin, family members' immigration statuses, and the mortality rates and trends in countries where family members reside abroad. If immigrants in the United States have family members residing in countries with higher mortality rates, this exclusion may result in an underestimation of population bereavement in the United States. In addition, this analysis follows Alburez-Gutierrez, Acosta, Zagheni, and Williams [44] in assuming independence between deaths, not accounting for clustering of losses within families. This simplifying assumption is supported by the same study, as [44] found that you would need very high levels of clustering (i.e., the ones found in high-intensity wars) for clustering to make a significant impact on bereavement estimates. In addition, the focus in this paper is on counting those experiencing bereavement annually, or in the case of the counterfactual analysis, those who might be spared from experiencing the loss of at least one close kin of different types during a given time. This is a conservative approach that does not capture reductions in the intensity of loss for those losing multiple siblings or children. While such reductions are undoubtedly important to surviving individuals and families, they have less of an impact on identifying individuals in need of psychosocial support for bereavement than capturing the change from never to ever bereaved. Also, this national-level analysis does not capture sub-national heterogeneity by factors such as race/ethnicity, education, or region, thereby potentially masking more severe bereavement in certain groups.

Our analysis builds on others that have examined population patterns in bereavement due to specific causes of death, specific types of familial loss, or across different social groups, and answers the call for more research engaging with the conceptualization of bereavement as a population-level experience [16]. Calculating trends in bereavement offers a new lens to the U.S. mortality crisis and highlights the public health problem of supporting bereaved individuals in the United States.

Materials and Methods

We applied a two-sex time-varying kinship model [45] to estimate the expected number of surviving kin for an average individual in the population, referred to as the "focal" individual (or Focal). Age- and sex-specific mortality rates and populations came from the Human Mortality Database [46], and age- and sex-specific fertility rates came from the Human Fertility Database [47]. Appendix A details the methodology and validation of imputing missing male fertility rates in specific years using the established relationship between male and female fertility while accounting for observed differences in the age distribution. The estimates of surviving kin were stratified by the age and sex of Focal and their kin. Kinship dynamics were modeled using a matrix projection framework, where each type of kin is treated as a separate population. Results from such kinship models have been validated against empirical data in European countries [48, 49]. More details on the kinship models can be found in Appendix B.1. All computations were conducted using the DemoKin R package [50].

To quantify bereavement, we followed the approach proposed by Acosta, Alburez-Gutierrez, Gargiulo, and Torres [36] for estimating bereaved persons. We used the kinship estimates and annual age- and sex-specific mortality rates to calculate the probability that a Focal at a specified age and sex lost at least one kin of a specified category (e.g., mother, father, sibling, or child). We then multiplied this probability by the at-risk population to estimate the number of bereaved individuals. Given evidence of the special importance of premature loss (defined as either death occurring at a young age or grieving at a young age), we pay special attention to losing a parent in ages 0–18 [18]. To plot trends in close kin losses at critical ages, we aggregated bereavement across sexes, broader age groups, and kinship types (e.g., females ages 0–18 who lost a parent). Details on bereavement calculations can be found in Appendix B.2.

Building on previous research [51, 52], we used counterfactual mortality scenarios to capture how many losses would not have occurred if the United States had maintained the rate of mortality improvement (ROMI) by age from 2000–2009 or if the United States had the ROMI of peer nations – Italy, Japan, Switzerland, and the United Kingdom. We calculated the ROMI for each country in two key periods: the pre-pandemic years (2010–2019) and the pandemic years (2019–2021) and then applied the period-wise ROMI to the observed U.S. mortality in 2009. Details on counterfactual analyses can be found in Appendix A.3.

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Appendix A: Data and Counterfactual Scenarios

A.1 Data Sources

Age- and sex-specific mortality rates and population exposures for each year and country were obtained from the Human Mortality Database [46]. Similarly, age-specific fertility rates (ASFR) for both females and males were retrieved from the Human Fertility Database [47].

A.2 Imputation of Missing Male Fertility Rates

For the United States, we have female fertility rates (ASFR_f)) available from 1933 to 2021. However, male fertility rates (ASFR_m) were missing for some years (before 1969 and after 2015). To address this, we imputed the missing values using the relationship between male and female fertility while accounting for observed differences in age distributions. The imputation procedure consists of the following steps.

A.2.1 Dynamic Age Adjustment

Male fertility generally peaks at older ages than female fertility. Traditionally, this difference is accounted for using a static age offset. Instead, we integrate a *dynamic* age difference term (ΔAge) into the regression model as a time-dependent variable. This approach allows the model to capture temporal shifts in male-female fertility relationships and better incorporate uncertainty. Figure A1 shows the age difference we calculated using the Human Fertility Database for the United States.

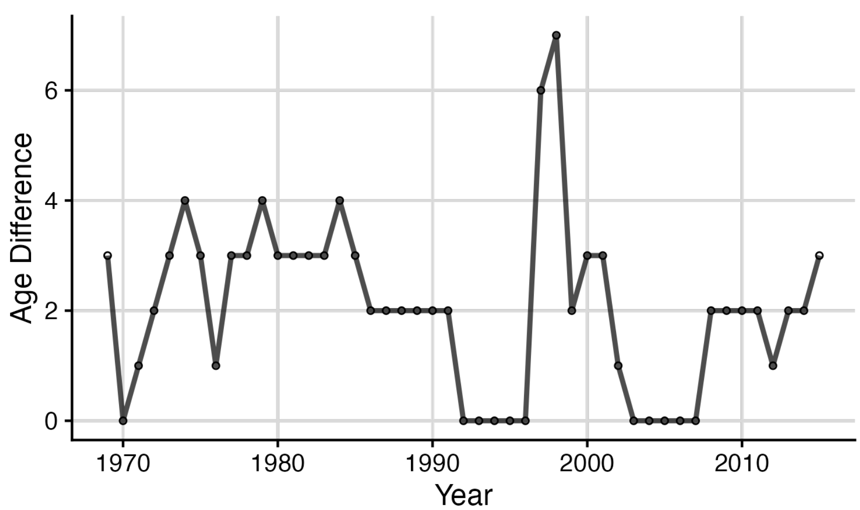

Figure A1: Sex difference between the peak age of age-specific fertility for the United States (Male - Female), from 1969 to 2015.

A.2.2 Regression Model

We model the male fertility rate (ASFR_m) as a function of female fertility (ASFR_f), age (Age), year (Year), and the time-varying age difference (ΔAge). The model specification is:

$$\begin{aligned}\text{ASFR}_m = \beta_0 + \beta_1 \times \text{ASFR}_f + \beta_2 \cdot \text{Age} + \beta_3 \times \text{Age}^2 + \beta_4 \times \text{Year} \\ + \beta_5 \times (\text{ASFR}_f \cdot \text{Age}) + \beta_6 \times \Delta\text{Age} + \beta_7 \times (\Delta\text{Age} \times \text{Year}) + \epsilon\end{aligned} \tag{1}$$

where:

- ASFR_f: Female age-specific fertility rate.
- Age: Age group (e.g., 5-year bins).
- Year: Calendar year of the observation.
- ΔAge: Difference between male and female peak fertility ages.

To account for nonlinearities and time-varying effects, we include interaction terms:

- $\text{ASFR}_f \times \text{Age}$ captures nonlinear interactions between female fertility rates and age.
- $\Delta\text{Age} \times \text{Year}$ allows for temporal variation in the male-female fertility offset.

For years and age groups with missing ASFR_m, we used the regression model to predict male fertility rates based on observed female fertility (ASFR_f), age, year, and ΔAge. Since there is no information on ΔAge when male fertility is missing, we assume $\Delta\text{Age} = 2$, which is approximately the mean value in Figure A1. In addition, the imputed male fertility has a few negative values at the young (< 18) and old (> 50) ages. Their magnitudes are trivial—the largest absolute value is less than 0.015—so we set these small negatives to zero.

A.2.3 Validation and Results

To validate the imputation method, we first compare the predicted and observed male fertility rates for years where male fertility records are available. Figure A2 presents this comparison across multiple decades, showing that the imputed values (dashed blue lines) successfully replicate key trends in male fertility (solid green lines), especially during 1980 and 2010. This includes the delayed peak relative to female fertility (solid red lines) and a broader age distribution. The strong alignment between observed and imputed values demonstrates the model's ability to capture sex-specific fertility differences.

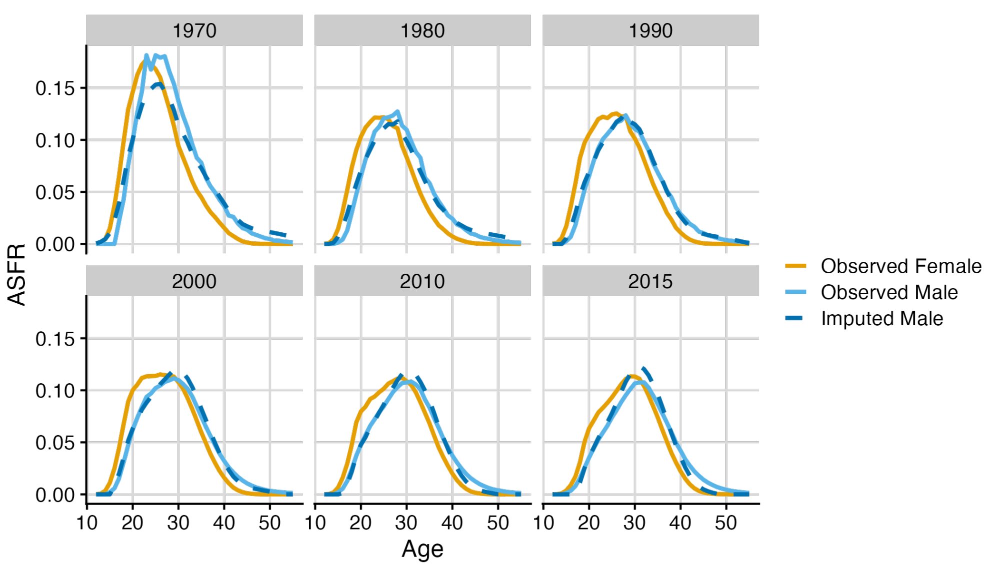

Figure A2: Comparison of observed and imputed male fertility rates, and observed female fertility rates across different years.

After validating the model with observed male fertility data, we apply the imputation to years where male fertility rates are missing. To assess the estimates, Figure A3 compares the imputed male fertility rates with observed female fertility rates. The results show that the model successfully extends the male fertility series while maintaining the expected male-female fertility offset and capturing shifts in fertility patterns over time.

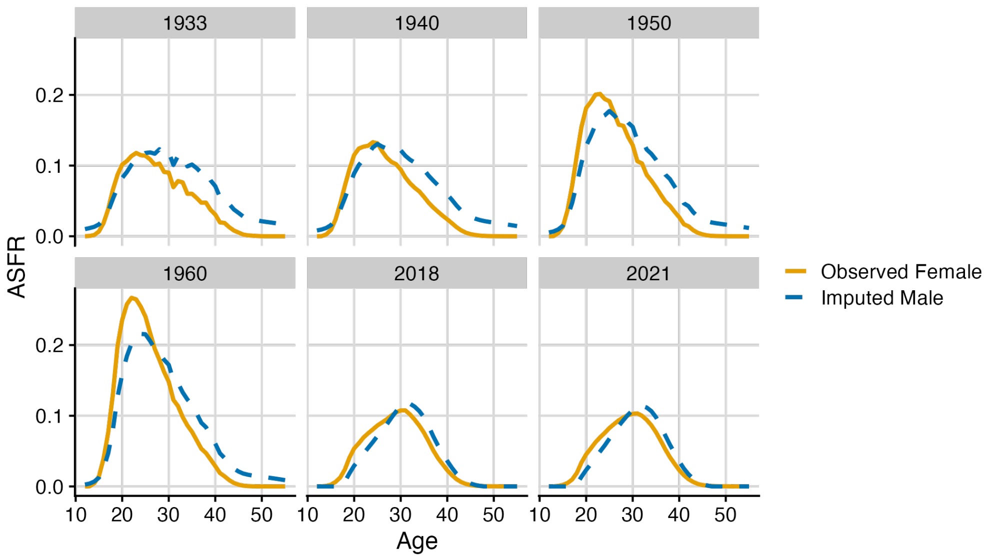

Figure A3: Comparison of observed female fertility rates and imputed male fertility rates across age in the United States

A.3 Counterfactual Mortality Scenarios

Building on previous research [51, 52], we use a simplified framework to examine mortality patterns in the United States from 2010 to 2021. Our counterfactual scenarios focus on international comparisons. We compare U.S. mortality trends with those of Italy, Japan, Switzerland, and the United Kingdom across all ages. To account for changes over time, we distinguish between two key periods: the pre-pandemic years (2010–2019) and the pandemic years (2019–2021). We calculate the Rate of Mortality Improvement (ROMI) for each country during each of the two periods. We then apply the period-wise ROMI to the observed US mortality in 2009 to construct counterfactual mortality scenarios after 2009. Additionally, we include a scenario where we apply the historical U.S. ROMI from 2000 to 2009 to project what U.S. mortality might have looked like had earlier trends continued.

Figure A4 shows the life expectancy at birth in the United States under different counterfactual scenarios for both females and males.

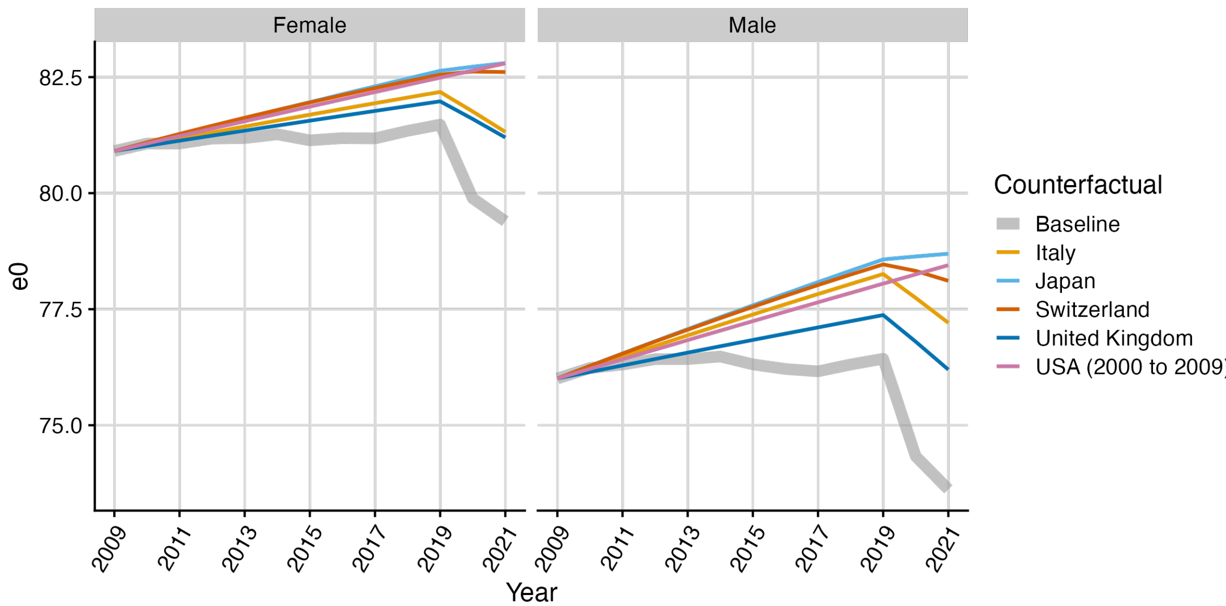

Figure A4: Life expectancy at birth (e_0) across different counterfactual scenarios, 2009 to 2021. Colors show different age scenarios, columns show different international comparisons, and rows show different sexes. The gray line indicates the baseline model with observed mortality.

Because the kinship model requires age-specific population exposures, we project populations under each mortality scenario (including the baseline) using the cohort component method [53]. These projections incorporate age-specific fertility and mortality rates while excluding migration effects.

Appendix B: Methodology of Bereavement Analysis

B.1 Two-Sex Time-Varying Kinship Model

We apply a two-sex time-varying kinship model [45] to estimate the expected number of surviving kin for an average individual in the population, referred to as the "focal" individual (or Focal). The estimates are stratified by both the sex and age of Focal and their kin. Kinship dynamics are modeled using a matrix projection framework, where each type of kin is treated as a separate population. All computations were conducted using the DemoKin R package [50].

B.1.1 Key Notation and Definitions

- $\mathbf{k}_f(x,t), \mathbf{k}_m(x,t)$: Age distribution for female and male living kin for each type of kin when Focal is at age x at time t.
- $\mathbf{U}_f(t), \mathbf{U}_m(t)$: Period female and male survival matrices at time t. Survival matrices contain age-specific survival probabilities on the subdiagonal.
- $\mathbf{F}_f(t), \mathbf{F}_m(t)$: Period female and male fertility matrices at time t. Fertility matrices include children of both female and male parents at the first row.
- $\alpha = 0.51$: Proportion of males among offspring.

B.1.2 Kinship Dynamics Projection

This is a summary of the approach presented in Caswell [45]'s work. For a more detailed derivation and comprehensive discussion of these models, readers should refer to that paper.

Based on the key notation, we first build the vector for kin estimate by combining male and female kin as:

$$\tilde{\mathbf{k}}(x,t) = \begin{pmatrix} \mathbf{k}_f \\ \mathbf{k}_m \end{pmatrix} (x,t), \tag{2}$$

Then we build the block matrices for the purpose of projection:

$$\tilde{\mathbf{U}}_t = \begin{pmatrix} \mathbf{U}_f(t) & 0 \\ 0 & \mathbf{U}_m(t) \end{pmatrix}, \quad \tilde{\mathbf{F}}_t = \begin{pmatrix} (1-\alpha)\mathbf{F}_f(t) & (1-\alpha)\mathbf{F}_m(t) \\ \alpha\mathbf{F}_f(t) & \alpha\mathbf{F}_m(t) \end{pmatrix}, \quad \tilde{\mathbf{F}}^*_t = \begin{pmatrix} (1-\alpha)\mathbf{F}_f(t) & 0 \\ \alpha\mathbf{F}_f(t) & 0 \end{pmatrix} \tag{3}$$

Finally, we can project the kinship dynamic as:

$$\tilde{\mathbf{k}}(x+1,t+1) = \tilde{\mathbf{U}}_t\tilde{\mathbf{k}}(x,t) + \tilde{\beta}(x,t), \tag{4}$$

where $\tilde{\beta}(x,t)$ represents the recruitment of new kin.

The two-sex model allows us to estimate kinship networks by both age and sex of the focal individual. In this way, $\tilde{\mathbf{k}}(x+1,t+1)$ is calculated for Focal of different sexes separately.

B.1.3 Kin Recruitment Scenarios

We consider three scenarios for kin recruitment $\tilde{\beta}(x,t)$.

The first scenario involves recruitment from the direct ancestors of the focal individual (e.g., parents, grandparents). This applies to kin such as younger siblings, who are derived from their direct parental lineage:

$$\tilde{\beta}(x,t) = \tilde{\mathbf{F}}^*\tilde{\mathbf{k}}^*(x,t), \tag{5}$$

where $\tilde{\mathbf{k}}^*(x,t)$ represents the source kin (e.g., the parents of the focal individual serve as the source kin for younger siblings of the Focal). The structure of $\tilde{\mathbf{k}}^*(x,t)$ is identical to that of $\tilde{\mathbf{k}}(x,t)$ used in the projection, with the asterisk (*) indicating that it refers to a different type of kin (source kin) rather than the kin being projected. In this case, we only use female fertility ($\tilde{\mathbf{F}}^*$) due to the interdependence of reproduction among pairs of ancestors. For instance, the mother and father of the focal individual do not reproduce independently to produce their offspring (i.e., the Focal's siblings). Consequently, we attribute reproduction to the mother, following the female fertility schedule.

The second scenario involves recruitment from other types of kin. This applies to kin such as cousins and nieces/nephews, who are recruited from the Focal's aunts/uncles and siblings, respectively:

$$\tilde{\beta}(x,t) = \tilde{\mathbf{F}}\tilde{\mathbf{k}}^*(x,t). \tag{6}$$

The third scenario assumes no new kin recruitment. This applies to kin who were born before the focal individual, such as parents, grandparents, and older siblings:

$$\tilde{\beta} = 0. \tag{7}$$

B.2 Bereavement Measurements

In this study, we quantify bereavement as the experience of losing at least one kin. We calculate two primary sets of metrics: the probability of an individual experiencing such a loss, and the total number of individuals in the population who are bereaved under this definition. This is achieved by moving from single-year, single-age probabilities to cumulative and aggregated metrics.

B.2.1 Notation

Table B1 summarizes the key variables and functions used in the bereavement calculations.

B.2.2 Probability of Losing Kin in a Single Year

We first establish the foundational, single-year probability of an individual experiencing bereavement.

We begin by defining the baseline probability that a Focal individual of age x_1 and sex s_1 does **not** lose any kin of type i in year t. We denote this kin preservation probability as π:

$$\pi(x_1,s_1,t,i) = \prod_{s_2}\prod_{x_2}(1-q(x_2,s_2,t))^{k(x_1,s_1,x_2,s_2,t,i)}. \tag{8}$$

Here, $k(x_1,s_1,x_2,s_2,t,i)$ is the expected number of kin for a Focal individual, obtained from the kinship model output $\tilde{\mathbf{k}}(x_1,t)$. The kin type i can represent a single category (e.g., "mother") or a collection of categories. For instance, in our analysis, we define a composite group **"close kin"** where i includes mother, father, siblings, and children. In such a case, we first aggregate the estimates for the various types of kin within a composition group. This yields an average number of all "close kin" (i.e., $k(x_1,s_1,x_2,s_2,t,i)$ where i includes multiple types of kin), then we proceed with the calculations starting from Equation 8.

Thus, the probability that the focal individual **experiences the loss of at least one kin** of type i in year t is the complement, λ:

$$\lambda(x_1,s_1,t,i) = 1-\pi(x_1,s_1,t,i). \tag{9}$$

Notation	Definition
Basic Parameters	
x_1, s_1	Age and sex of the focal individual. Age x_1 is typically anchored to a specific time point (e.g., t, t_1 or t_2).
x_2, s_2	Age and sex of the kin relative.
i	Kin category (e.g., parents, siblings, etc.).
t, t_1, t_2	Time point (year), start of an interval, and end of an interval.
A	A set representing an age group for the focal individual.
$q(x_2, s_2, t)$	Probability of dying for of individuals aged x_2, sex s_2 at time t.
$k(x_1, s_1, x_2, s_2, t, i)$	Expected number of kin of type i when the focal individual is aged x_1 with sex s_1, and kin is aged x_2 with sex s_2 at time t
$\mathrm{pop}(x_1, s_1, t)$	Population exposure of individuals aged x_1, sex s_1 at time t.
Bereavement Functions (Loss of at least one kin)	
$\pi(x_1, s_1, t, i)$	Probability of **not** losing any kin of type i for a Focal aged x_1 with sex s_1 in a single year t.
$\lambda(x_1, s_1, t, i)$	Probability of **experiencing the loss of at least one kin** of type i for a Focal aged x_1 with sex s_1 in year t.
$\Pi(x_1, s_1, t_1, t_2, i)$	Probability of **not** losing any kin of type i over $[t_1, t_2]$, for a Focal who is age x_1 with sex s_1 at time t_2.
$\Lambda(x_1, s_1, t_1, t_2, i)$	Cumulative probability of **experiencing at least one kin loss** of type i over $[t_1, t_2]$, for a Focal who is age x_1 with sex s_1 at time t_2.
$B(x_1, s_1, t_1, t_2, i)$	Number of individuals (age x_1 at t_2) with sex s_1 who have experienced at least one kin loss over $[t_1, t_2]$. When $t_1 = t_2 = t$, it is a single-year estimation of the **the number of individuals experiencing at least one kin loss in a specific year** t. When $t1 < t_2$, it is a period estimation of the **cumulative number of ever-bereaved individuals during a time period**.

Table B1: Summary of notation used in bereavement estimation. The functions p, q, B can operate on single years or time intervals, with age correctly indexed over time.

B.2.3 Probabilities over Time Intervals

We then generalize the age- and year-specific probabilities to capture the cumulative experience of bereavement over any time interval $[t_1, t_2]$. We note that an individual's age changes with time and we define our interval functions based on the focal individual's age x_1 at the **end** of the period, t_2.

The probability that a focal individual (who is age x_1 at t_2) does **not** lose any kin throughout the entire period from t_1 to t_2 is the product of the age- and year-specific survival probabilities, where the age of the Focal is dynamically updated each year. At any given year τ in the interval, the Focal's age is $x_1 - (t_2 - \tau)$.

$$\Pi(x_1, s_1, t_1, t_2, i) = \prod_{\tau=t_1}^{t_2} \pi(x_1 - (t_2 - \tau), s_1, \tau, k). \tag{10}$$

This generalized notation also covers the single-year estimation as $\Pi(x_1, s_1, t_1, t_2, i) = \pi(x_1 - (t_2 - \tau), s_1, \tau, k)$ when $t_1 = t_2 = \tau$.

Consequently, the cumulative probability that a focal individual has **experienced the loss of at least one kin** at any point during the interval $[t_1, t_2]$ (i.e., the probability of being "ever bereaved" during this period) is:

$$\Lambda(x_1, s_1, t_1, t_2, i) = 1 - \Pi(x_1, s_1, t_1, t_2, i). \tag{11}$$

Similarly, this generalized notation also covers both single-age and multi-year probabilities, where $\lambda(x_1, s_1, t, i) = \Lambda(x_1, s_1, t, t, i)$.

B.2.4 Number of Annual Bereaved Persons and Cumulative Number of Ever-Bereaved Persons

Next, we translate probabilities into population-level counts of individuals who have experienced kin loss.

We define a general function $B(x_1, s_1, t_1, t_2, i)$ for the number of individuals bereaved. The number of individuals experiencing at least one kin loss **in a specific year** t, which is the annual number of kin loss incidents, is:

$$B(x_1, s_1, t, t, i) = \Lambda(x_1, s_1, t, t, i) \times \mathrm{pop}(x_1, s_1, t). \tag{12}$$

Here, the beginning and end years are both year t since we are doing single-year estimations. Similarly, the number of individuals alive at t_2 with age x_1 who have experienced at least one kin loss since t_1, which is the cumulative number of ever-bereaved individuals **during a time period**, is:

$$B(x_1, s_1, t_1, t_2, i) = \Lambda(x_1, s_1, t_1, t_2, i) \times \mathrm{pop}(x_1, s_1, t_2). \tag{13}$$

B.2.5 Aggregated Measurements Across Ages

We also scale our metrics from individuals at each single age to entire age groups to measure the population-wide burden of bereavement. To analyze bereavement for a broader age group A, we define aggregated metrics for both single-year incidence and multi-year cumulative prevalence of experiencing at least one kin loss.

For a single year, the aggregated annual number of bereaved individuals in age group A in year t (beginning with year t and ending with year t) is the population-weighted average:

$$\Lambda_{aggr}(A, s_1, t, t, i) = \frac{\sum_{x_1 \in A} B(x_1, s_1, t, t, i)}{\sum_{x_1 \in A} \mathrm{pop}(x_1, s_1, t)}. \tag{14}$$

The total number of annual bereaved individuals in age group A in year t is the sum:

$$B_{aggr}(A, s_1, t, t, i) = \sum_{x_1 \in A} B(x_1, s_1, t, t, i). \tag{15}$$

For a time period, the aggregated proportion of the population in age group A who have been **ever bereaved** by year t_2 since t_1 is:

$$\Lambda_{aggr}(A, s_1, t_1, t_2, i) = \frac{\sum_{x_1 \in A} B(x_1, s_1, t_1, t_2, i)}{\sum_{x_1 \in A} \text{pop}(x_1, s_1, t_2)}. \tag{16}$$

And the total number of individuals in age group A who have experienced at least one kin loss over the interval is:

$$B_{aggr}(A, s_1, t_1, t_2, i) = \sum_{x_1 \in A} B(x_1, s_1, t_1, t_2, i). \tag{17}$$

B.2.6 Decomposition of Bereaved Persons

Finally, we attribute differences in the annual number of persons who experienced kin loss in a specific year between populations to key demographic components. To understand why the annual number of individuals experiencing bereavement differs between two populations (indexed 1 and 2), we can apply a Kitagawa-Keyfitz decomposition to the single-year, single-age counts from Eq. 12. For simplicity, we remove the function arguments (x_1, s_1, t, i):

$$B_1 - B_2 = (\Lambda_1 - \Lambda_2) \times \frac{\text{pop}_1 + \text{pop}_2}{2} + (\text{pop}_1 - \text{pop}_2) \times \frac{\Lambda_1 + \Lambda_2}{2}. \tag{18}$$

The first term on the right-hand side captures the effect of differing probabilities of kin loss (the rate effect), while the second term captures the influence of population size and age structure (the composition effect). This decomposition can first be applied at the granular level of x_1, s_1, t, and i and then aggregated as needed for the study (e.g., across ages, years, kin types, etc.).

Appendix C: Figures

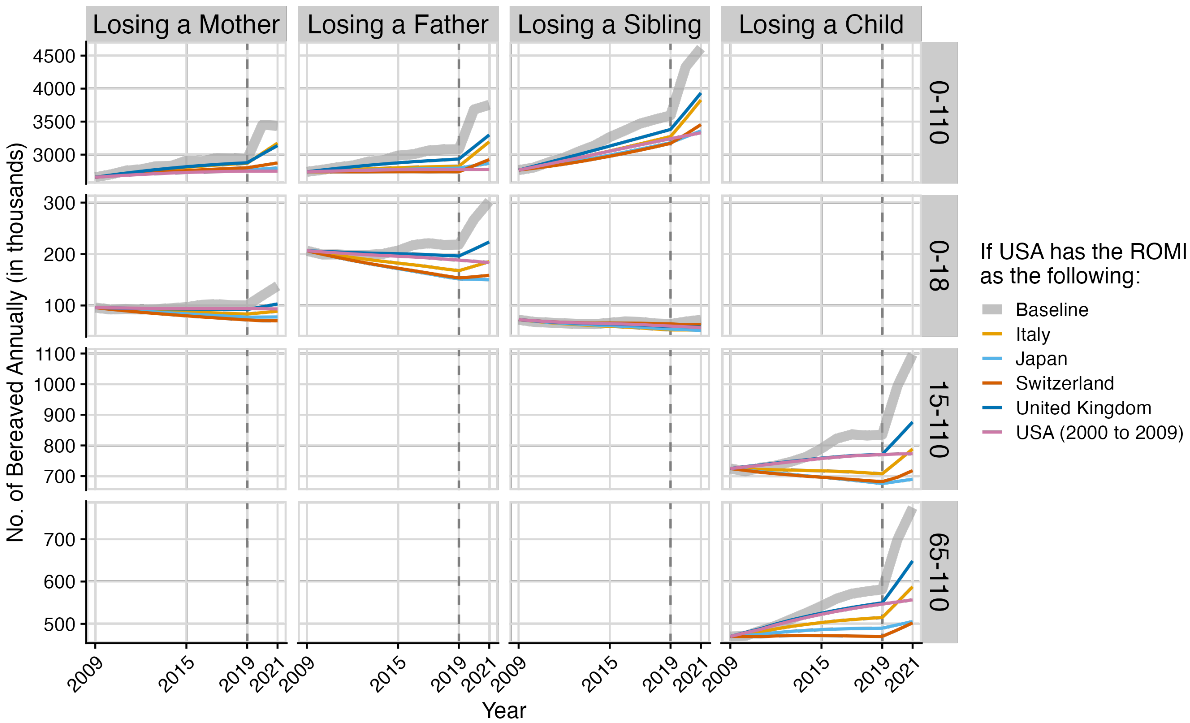

Figure C1: Number of individuals experiencing kin loss annually from 2009 to 2021, comparing baseline estimates (solid gray lines) with counterfactual scenarios (indicated by different colors). Here, counterfactual scenarios assume that the United States had the rate of mortality improvement (ROMI) of peer nations in 2010–2021, or it had maintained its rate of mortality improvement from 2000–2009 in 2010–2021. Colors represent the difference between the baseline and different counterfactual scenarios. Results are shown by bereaved individuals of different age groups (by rows) and by the different types of kin they lose (by columns). Dashed gray line marks 2019.

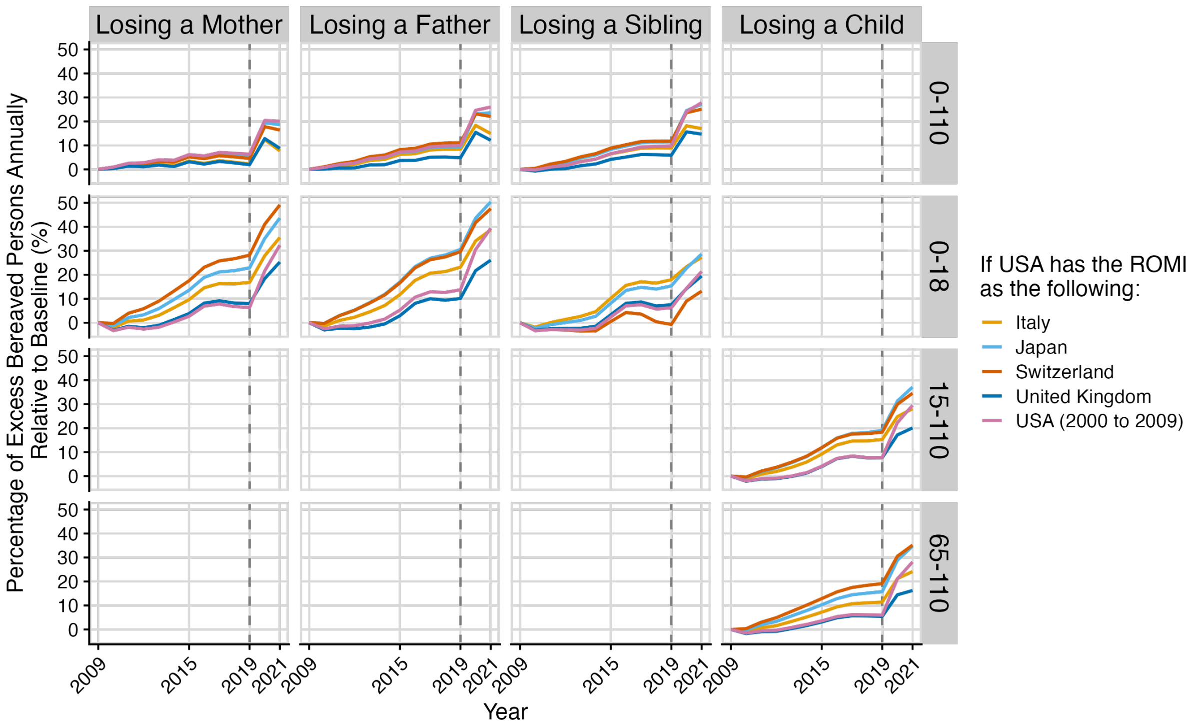

Figure C2: Percentage of excess bereaved persons annually from 2009 to 2021, calculated as bereaved persons in baseline minus each counterfactual scenario divided by the baseline. Here, counterfactual scenarios assume that the United States had the rate of mortality improvement (ROMI) of peer nations in 2010–2021, or it had maintained its rate of mortality improvement from 2000–2009 in 2010–2021. Colors represent the percentage difference between the baseline and different counterfactual scenarios. Results are shown by bereaved individuals of different age groups (by rows) and by the different types of kin they lose (by columns). Dashed gray line marks 2019.

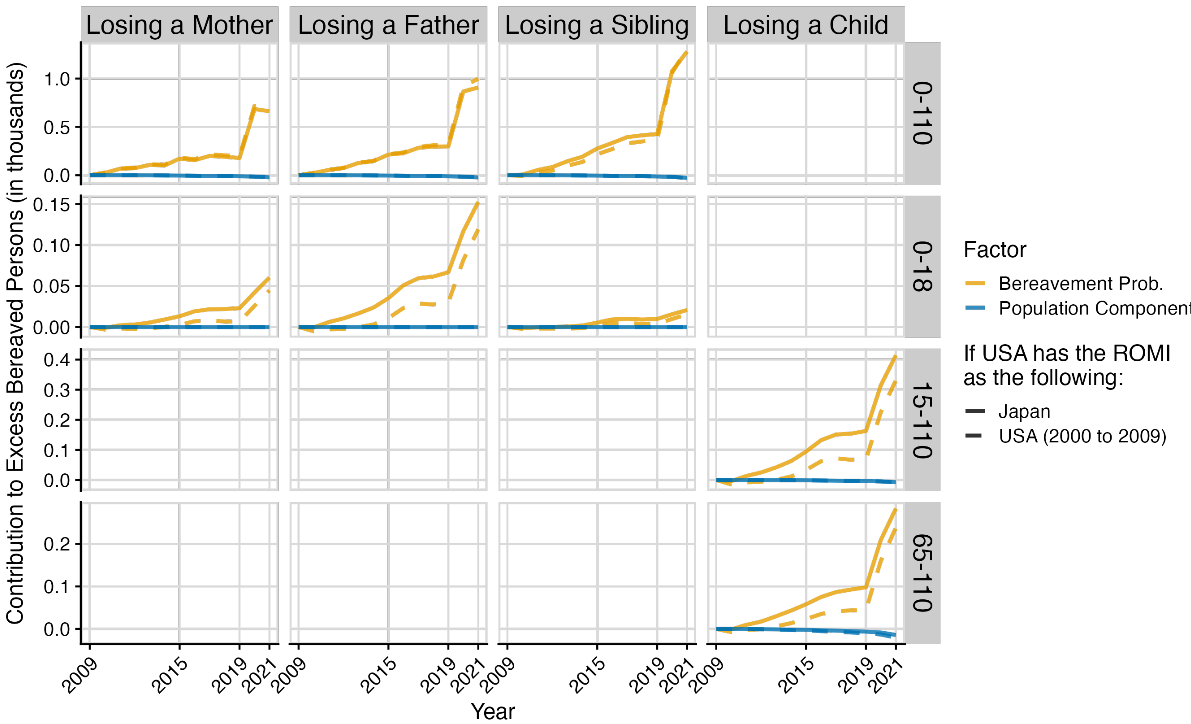

Figure C3: Contributions of bereavement probability and population component to the excess bereaved persons across years (in thousands). Such excess bereaved persons are calculated as the difference in the number of bereaved persons between baseline and counterfactual scenarios. Here, counterfactual scenarios assume that the United States had the rate of mortality improvement (ROMI) of peer nations in 2010–2021, or it had maintained its rate of mortality improvement from 2000–2009 in 2010–2021. Colors represent the contribution of different factors. Lines show different counterfactual scenarios. Results are shown by bereaved individuals of different age groups (by rows) and by the different types of kin they lose (by columns).