NBER WORKING PAPER SERIES

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Working Paper 26638
http://www.nber.org/papers/w26638

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
January 2020

We are thankful to David Card, Janet Currie, Bill Evans, Adriana Lleras-Muney, Steve Pischke, Julian Reif, Jonathan Skinner, Ann Stevens, seminar participants at UC Davis, Cemfi Madrid, Aarhus, Hohenheim, Ifo, University of Connecticut, and conference participants at NBER Summer Institute 2019, NBER Labor Studies meeting 2016, NBER Universities' Research Conference 2013, Federal Reserve Bank of Minneapolis, Barcelona GSE Summer Forum, Society of Labor Economics 2019, International Health Economics Associa- tion 2019, and the American Society of Health Economics 2019. Nicolas Oderbolz provided excellent research assistance. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

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NBER Working Paper No. 26638
January 2020
JEL No. E32,I10,J10

ABSTRACT

This paper uses several large cross-sectional data sources and a new approach to estimate midlife effects of entering the labor market in a recession on mortality by cause and various measures of socioeconomic status. We find that cohorts coming of age during the deep recession of the early 1980s suffer increases in mortality that appear in their late 30s and further strengthen through age 50. We show these mortality impacts are driven by disease-related causes such as heart disease, lung cancer, and liver disease, as well as drug overdoses. At the same time, unlucky middle-aged labor market entrants earn less and work more while receiving less welfare support. They are also less likely to be married, more likely to be divorced, and experience higher rates of childlessness. Our findings demonstrate that tempo-rary disadvantages in the labor market during young adulthood can have substantial impacts on lifetime outcomes, can affect life and death in middle age, and go beyond the transitory initial career effects typically studied.

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1 Introduction

Economists and policy makers have long been concerned that entering the labor market in a recession could worsen lifetime outcomes of young workers through no fault of their own.[2] Research has indeed shown that entering the labor market in a recession leads to large initial losses in earnings and employment. These studies typically find the initial effects fade after about ten years in the labor market (e.g., Oyer, 2006; Kahn, 2010; Wozniak, 2010; Oreopoulos et al., 2012; Schwandt and von Wachter, 2019; Rothstein, 2019). Less is currently known about the longer-term effects of adverse labor market entry as these cohorts go through middle age, or how these shocks affect outcomes beyond earnings that become more relevant in middle age, such as mortality and broader measures of socioeconomic status (SES).

Yet, some of the available evidence suggests that the initial effects could linger past the ten to fifteen years usually studied. Losses in cumulated lifetime income implied by typical estimates per se could lead to lower wealth accumulation, and there is some evidence of reductions in housing wealth among individuals coming of age in the Great Recession (e.g., Dettling and Hsu, 2014). Several studies have documented lasting changes in occupational choice (Oyer, 2006, 2008; Altonji et al., 2016) and employer characteristics (Oreopoulos et al., 2012), and Kahn (2010) finds that 1982 college graduates may begin to lose ground again around 15 years after job entry.

Moreover, unlucky labor market entry may have impacts beyond employment and earnings that affect SES more broadly and worsen longer-term health. For example, Currie and Schwandt (2014) show that recession exposure in early adulthood reduces lifetime fertility for women. There is also evidence that male college graduates from the 1982 recession experience worsening self-reported health in middle age (Maclean, 2013). Potential longer-term health effects may be a persistent consequence of initial declines in career outcomes and ensuing effects on socioeconomic status. In addition, evidence suggests that the period of school-leaving can be a formative age for behavioral health (Kessler et al., 2005). Overall, these findings raise concerns that entering the labor market in a recession may have larger consequences than current estimates on earnings suggest. Yet, so far these broader and longer-term effects have not been the focus of the literature on recession entrants.

[2] See, e.g., New York Times, 2006, "Hello Young Workers: One Way to Reach the Top is To Start There" ;Wall Street Journal, 2009, "The Curse of the Class of 2009"; Hearing of the Joint Committee of Congress, 2010, "Avoiding a Lost Generation: How to Minimize the Impact of the Great Recession on Young Workers"; The Economist, 2018, "Temporary economic downturns have long-lasting consequences"; Wall Street Journal, 2018, "Playing Catch-Up in the Game of Life,"; Financial Times, 2018, "The untold career value of a little bit of luck at the outset."

In this paper, we estimate the midlife effects of graduating during a recession on mortality and various measures of socioeconomic status. Using several large cross-sectional data sources and a new approach addressing data limitations and potential selection bias, we find that cohorts that entered the labor market during the early 1980s recession experience increases in mortality in middle age. The increases in mortality are chiefly driven by certain causes of death, including heart disease, lung cancer, liver disease, and drug poisoning. We also show that these cohorts experience a rise in divorce rates, have fewer children, and experience a widening gap in earnings with respect to luckier labor market entrants.

We focus on mortality because it is an important and standard measure of population health that is reliably measured for a large number of cohorts throughout the life cycle. Measuring the long-term effect of adverse labor market entry on mortality is also important in light of the ongoing debate on the nexus between economic outcomes and health and longevity. Despite much work on the topic, the direction of causality remains controversial and careful research suggests that indicators of economic prosperity, such as income or wealth, may have little measurable causal effects on health (for a review, see Cutler et al. (2011a)).

One example is the literature on business cycles and health, pioneered by Ruhm (2000), which finds that aggregate mortality is *lower* during times of high unemployment. Another example is the discussion surrounding the sources of recent increases in midlife mortality related to drug overdoses, liver disease, and suicide uncovered by Case and Deaton (2015). Case and Deaton (2015, 2017, 2018) argue that socioeconomic hardship is a major driving force fueling what they coin a "deaths of despair" crisis. But despite a growing number of studies analyzing the relationship of economic conditions and increasing midlife mortality, it remains debated whether a causal relationship can be detected in data (Carpenter et al., 2017; Ruhm, 2018; Currie et al., 2018; Bound et al., 2018; Ruhm, 2019).[3]

Our study addresses one particular causal channel between economic conditions and mortality, running from initial labor market conditions to long-term mortality outcomes of affected cohorts in middle age. A key advantage of our approach is that aggregate economic fluctuations such as the unemployment rate are unlikely to be driven by individual health differences. Moreover, by focusing on labor market entrants, we analyze a group that experiences a large economic shock that is unrelated to their own earnings potential. Finally, by focusing chiefly on the longer-term effects, we deal with the fact that health responses to economic conditions may take time to accumulate and therefore might be hard to detect in

[3]For recent media articles on that debate see NewYorkTimes (2019) and WashingtonPost (2019).

a more contemporaneous analysis. To better understand the context of our mortality results, we also exploit our methodology to examine the effect on a range of SES measures in middle age. Studying these outcomes is interesting in its own right because they may be directly relevant beyond earnings and because they may reflect adjustment costs that individuals may have incurred to rebuild their careers and to recover initial lost earnings. These outcomes are also relevant in terms of potential relationships to mortality stressed by prior literature.

Our approach focuses on fluctuations in the state-level unemployment rate, which provides us with exogenous variation while allowing us to account for potentially confounding trends at the national level. Hence, our analysis is not able to capture effects of national economic conditions on labor market entrants. To analyze effects in middle age, we focus on cohorts entering the labor market in different U.S. states before, during, and after the 1982 recession — the largest post-war recession before the Great Recession — that we can observe from labor market entry until at least age 49. We use Vital Statistics data from 1979 to 2016 and population estimates from the Census and the ACS to construct mortality rates that are regressed on the state-level unemployment rate that a cohort faced at the time of graduation. Information on socioeconomic outcomes, including earnings, labor supply, marital status, divorce, and cohabitation is derived from the Decennial Census (Census), the American Community Survey (ACS), and the Current Population Survey (CPS).

The large sample sizes of our cross-sectional data sources are important to capture the effects we seek to measure. One complication is that these data do not contain information on the state and the year in which an individual entered the labor market. Furthermore, people might migrate to a different state before graduating or adjust their graduation year in response to local economic conditions. These selection responses can severely bias the analysis even if we had a perfect measure of the location and the time of graduation. We address these measurement and selection issues using a method introduced by Schwandt and von Wachter (2019). As explained in the paper, we re-weight unemployment rates to reflect the economic conditions a cohort would face at graduation if it had migration and education rates similar to surrounding cohorts. A key advantage of this approach is that it only requires information at the birth-state and birth-year level and hence can be applied to Vital Statistics data, which otherwise could not be used to study the long-term effects of state-level labor market entry shocks.

Using this approach, we obtain four main findings. First, we find that a temporarily higher state unemployment rate at the age of labor market entry leads to precisely estimated increases in mortality in middle age. The magnitude of these effects is meaningful – if

sustained until the end of their life, a 3.9 percentage point higher unemployment rate (as experienced by the 1982 graduation cohort) would lead to a decrease in life expectancy by six to nine months.

Our second main findings is that these midlife mortality increases are driven to an important extent by increases in mortality due to specific causes of death. These causes are both disease-related and "external" causes, such as lung cancer, liver disease, and drug poisoning. Hence, entering the labor market during a recession raises mortality by increasing important elements of "deaths of despair" in the language of Case and Deaton (2015, 2017). In contrast, we find no long-term effects on other causes of death, such as accidents or cancers other than lung cancer.

Our third main finding is that entering the labor market during a recession has a substantial impact on a range of SES measures in middle age: unlucky cohorts show lower marriage rates, higher divorce rates, and smaller family size. We also find that after initial earnings recovery in their mid-thirties, adversely affected entry cohorts suffer a reduction in earnings as they reach their mid-forties, confirming Kahn's (2010) exploration of college graduates from the early 1980s recession based on a shorter time period. Our research design does not allow us to disentangle potentially complex causal channels between these outcomes and mortality. But these results make clear that entering the labor market during the 1980s recession had broad effects in middle age on a range of key life outcomes that may influence, and may itself be influenced by, underlying health related to mortality.

Our fourth main finding is that there is some heterogeneity in the effects of the initial unemployment rate on midlife outcomes. While the effects on overall mortality are similar by race, increases in our measure of "deaths of despair" appear to be concentrated among non-Hispanic whites. White men also tend to experience a decline in earnings in midlife, and tend to experience larger reductions in family stability than their non-white counterparts. This is despite the fact that non-whites experience larger short-run effects on earnings and other outcomes (Schwandt and von Wachter, 2019). An important exception is that non-whites experience a lasting reduction in receipt of welfare income, without an appreciable rise in earnings.

Our paper makes several contributions to related strands of literature. Ours is the first paper to study the effect of entering the labor market in a recession on mortality. So far, the literature has not focused on mortality likely because mortality rates at younger ages are small and the effects on mortality likely take longer to materialize. We are also the first to explicitly focus on the longer-term effects of adverse labor market entry.

The paper thereby extends a growing body of literature studying the short- to medium-term effects on key labor market outcomes such as earnings, wages, and employment (Oyer, 2006, 2008; Kahn, 2010; Wozniak, 2010; Genda et al., 2010; Oreopoulos et al., 2012; Altonji et al., 2016; Schwandt and von Wachter, 2019), and a smaller literature studying these effects on fertility and marital status (e.g., Kondo 2011; Currie and Schwandt 2014) and health (Maclean, 2013; Cutler et al., 2015).[4] The paper also relates to studies estimating the causal effects of specific labor market shocks on health and mortality at other ages, complementing previous studies that find persistent mortality effects for economic conditions around birth (Van den Berg et al., 2006), adolescence (Cutler et al., 2016), and before retirement (Coile et al., 2014).

With respect to papers looking at short- and long-term effects of other labor market shocks, such as job displacements, on mortality (e.g., Sullivan and von Wachter, 2009; Browning and Henesen, 2012; Eliason and Storrie, 2009) or health (e.g., Schaller and Stevens 2015; Kuhn et al. 2009; Burgard et al. 2007; Browning et al. 2006) that typically analyze workers who lose stable jobs due to a downsizing or plant closure closer to middle age, our paper studies the long-term effects of economic shocks on young labor market entrants.[5] In contrast to those papers, we study an aggregate shock, which is a less specific treatment, but which may be more clearly exogenous with respect to workers' own health or productivity.

Our paper is also related to the broader literature that seeks to establish a causal link between income or SES and mortality (e.g., Cutler et al., 2011b; Sullivan and von Wachter, 2009; Cutler et al., 2016; Schwandt, 2018). An advantage of our paper is that we exploit external labor market shocks to establish a causal link from economic conditions to mortality. A potential drawback of this approach is that we cannot obtain the causal effect of a specific measure of income or SES. However, our finding that a broad range of SES measures is affected by the economic shock suggests that it may be difficult to find a single summary measure for shifts in economic status. Instead, the mortality effect itself could be thought of as a summary measure of those long-term effects of economic shocks that individuals could not buffer through mechanisms available to them. Clearly, this does not capture all the costs of the shock to individuals, but it represents a particularly stark adverse consequence.

Within the broader literature on socioeconomic outcomes and mortality, our findings help to shed light on several phenomena discussed in the recent literature. We replicate

[4] Studies suggest other social outcomes that may be related to health worsen for cohorts entering the labor market in recessions as well, including increases in crime (e.g., Bell et al. 2018) and reductions in self-esteem (e.g., Maclean et al. 2016).

[5] A closely related literature studies the effect of individual unemployment spells on mortality and other health outcomes (e.g., Halliday, 2014; McLeod et al., 2012; Roelfs et al., 2011; Martikainen et al., 2007).

the reverse, negative relationship of mortality and contemporaneous unemployment rates for young labor market entrants (e.g., Ruhm, 2000; Miller et al., 2009; Stevens et al., 2015; Ruhm, 2015; Hollingsworth et al., 2017). But we find this relationship reverses sign and becomes positive in the long run for these cohorts. This suggests that a positive SES-health relationship may appear in mortality data only after a substantial lag. Increasing SES-health gradients with age have been well-documented in the empirical literature (e.g., Case et al., 2002; Currie and Stabile, 2003) and recently incorporated in a model of life-cycle mortality by Moreau and Lleras-Muney (2019). Yet, there have been doubts about how much of the gradient and its rise represent true causal effects of SES on health (e.g., Cutler et al., 2011a). Our results are consistent with the notion that initial differences in SES tend to have an increasing impact on health and eventually on mortality as individuals become older.

Moreover, our paper establishes a direct causal link between initial economic conditions, long-term socioeconomic decline, and mortality among cohorts at the core of recent increases in midlife mortality (Case and Deaton, 2015, 2017; Borella et al., 2019). This supports the notion brought forward by Case and Deaton (2015, 2017, 2018) that persistent social and economic hardship may be an important channel behind current midlife mortality crisis. However, it is important to point out that by design, our paper is not meant to explain the entire or even a large fraction of the midlife mortality crisis, since we look only at initial labor market conditions orthogonal to the national unemployment rate, and neither address the effects of recessions for older worker nor the role of wider economic trends, e.g. related to trade (Pierce and Schott, forthcoming; Autor et al., forthcoming).

Finally, by showing how the concept of the "double-weighted" unemployment rate introduced by Schwandt and von Wachter (2019) can be used to study initial conditions in large cross-sectional mortality data sets, our paper makes a useful conceptual contribution that we expect will benefit a range of future studies based on similar administrative data. Taking this step is important because large, administrative data spanning many cohorts and years is required to study the effect of long-term aggregate shocks on mortality – due to infrequently occurring shocks, moderate effects sizes, and substantially lagged effects as cohorts age – but typically vital statistics data lag key information on labor market history typically used for this kind of analysis.

The remainder of the paper proceeds as follows. The next section introduces the data while the third section lays out our estimation approach. The fourth section presents our main findings on the effects of entering the labor market in a recession on mortality in middle age, overall and by cause of death. The fifth section summarizes our findings on long-term

responses in measures of SES. The sixth section discusses our findings in light of common economic mechanisms. The final section concludes.

2 Data

Our analysis is based on four data sets: the U.S. Vital Statistics (Vital Stats), which provides information on every death in the United States since 1968; the Decennial Census of Population (Census) in combination with the American Community Survey (ACS), which provide the population denominator for the construction of annual mortality rates as well as socioeconomic characteristics; the Annual Social and Economic Supplement of the March Current Population Survey (CPS), which provides socioeconomic characteristics for repeated cross-sections for large samples of individuals; and annual state-level unemployment rates from the Bureau of Labor Statistics.

Annual mortality rates require annual population estimates. We construct population estimates by state of birth and year of birth using the Census years 1980, 1990, and 2000 and ACS annual estimates for 2001–2016. We apply a linear interpolation for the pre-2000 inter-census years. The death files from the Vital Stats include the cause of death along with core demographic characteristics such as age, sex, race, and state of birth. The same characteristics are reported in the Census/ACS, which allows us to construct death rates for the subgroups defined by these demographics.[6]

We analyze labor market and family formation outcomes from pooled data from the CPS (waves 1979–2016), the Census (1980, 1990, 2000), and the ACS (2001–2016). Our baseline specification focuses on U.S.-born individuals and links the unemployment rate at labor market entry via individuals' state and year of birth. The CPS does not contain the state of birth so we use the state of current residence as a proxy (see Schwandt and von Wachter, 2019, and Section 5.1 for a discussion). We leave the CPS data unrestricted in terms of nativity in years prior to 1994, as the country of birth is only reported afterwards. We analyze both mortality and socioeconomic outcomes in four demographic subgroups: male and female non-Hispanic whites (referred to as male and female whites), and male and female non-whites and Hispanics (referred to as male and female non-whites).

Our main economic outcome is log earnings, which refers an individual's total pre-tax

[6] After 1989, the mortality files additionally include information on the decedent's education; thus in principle, we could construct mortality rates by education group and graduation year for the years following 1989. However, constructing mortality rates by education group is problematic due to missing information and coding mismatches between the Vital Stats and the Census (Currie and Schwandt, 2016b).

wage and salary income in a given calendar year. Labor force participation is measured as reporting any positive earnings. Welfare receipt includes total annual pre-tax income received from public assistance programs.[7]Poverty status is assessed by comparing family income to the federal poverty line. Family formation measures include current marital status and the number of own children currently living in the respondent's household (the number of own children ever born is not reported in the CPS or Census/ACS data after 2000).

3 Empirical Approach

To obtain the causal effect of economic conditions at job market entry on mortality and socioeconomic outcomes we would ideally like to estimate the following regression model:

$$outcome_{i,t} = \alpha + \beta_a ec_{i0} + \gamma_a + \epsilon_{i,t} \tag{1}$$

where $outcome_{i,t}$ are different outcomes of an individual i at different ages a, ec_{i0} are local labor market conditions at labor market entry, and γ_a are age fixed effects. The coefficients β_a represent age-specific deviations from the typical age profile resulting from differences in initial local labor market conditions. To implement the equation, we have to choose a level aggregation for the appropriate labor market. For our analysis, we follow the majority of the literature and use annual state-level unemployment rates as indicators for the state of the local labor market. For a causal interpretation of β_a we further require that economic conditions at labor market entry be orthogonal to other determinants of the respective outcome that are not captured by the model.

3.1 Identification strategy

There are two main challenges to estimating Equation (1). The first challenge is a data limitation: the Vital Stats mortality data and other population-wide data sources do not contain the state and year of graduation. The second challenge is a selection issue: if people select when to graduate or where to graduate in response to local labor market conditions, then economic conditions at labor market entry are not orthogonal to other factors potentially impacting an individual's later outcomes.

We follow the method developed by Schwandt and von Wachter (2019) to address both

[7]These include Supplemental Security Income payments to blind or disabled persons with low incomes; Aid to Families with Dependent Children; and General Assistance (not including separate payments for hospital or other medical care).

challenges. We use data from the Census and the ACS to construct a proxy for the graduation year unemployment rate that only relies on information on an individual's state of birth and year of birth. Because these characteristics are fixed, they cannot be affected by labor market conditions around graduation. A key advantage of this approach is that these characteristics can be linked to the Vital Stats data, which also report the state and year of birth.[8]

To implement this approach, we first estimate migration shares $\overline{mig}^A_{b,s}$ as the average share of cohorts in our sample born in state b that live in state s at ages $A = 16, 18, 20, 22$. Note that we only use state-specific *average* migration rates (i.e., state fixed-effects in the underlying regression model of individual level migration indicators) rather than migration rates of a specific birth cohort (state-cohort fixed effects), which could be driven by an endogenous response to contemporaneous labor market conditions. Next, we estimate graduation shares $\overline{edu}^A_{b,c}$, indicating the share of sample cohorts born in state b in year c who graduate at age $A = 16, 18, 20, 22+$. We predict graduation shares using both state fixed effects and country-wide cohort fixed effects (but not their interaction). The "double-weighted" (DW) average graduation year unemployment rate is then given by:

$$u^{DW}_{b,c} = \sum_A \overline{edu}^A_{b,c} \sum_{s=1}^{51} \overline{mig}^A_{b,s} u_{s,c+A} \tag{2}$$

The double-weighted unemployment rate does not rely on endogenous migration at labor market entry or timing of labor market entry in response to local labor market conditions, as it is constructed using only state averages over time and national averages across periods, but not state-period variation. Moreover, because the cohorts are assigned their graduation state based on their state of birth, which is a fixed characteristic, the double-weighted unemployment rate is also not affected by state-cohort-specific migration after graduation.

3.2 Double-weighted specification (baseline)

We regress mortality and other outcomes on the double-weighted average unemployment rate, using the following specification:

$$outcome_{b,c,a} = \alpha + \beta_a u^{DW}_{b,c} + \gamma_a + \lambda_b + \delta_c + \theta_t + \epsilon_{b,c,a} \tag{3}$$

The indices b, c, a, and t refer to the birth state, birth year (cohort), age, and calendar

[8]This approach is based on synthetic cohorts (Deaton, 1985) and is further developed in Currie and Schwandt (2014), which link maternal life cycles to unemployment rates using mothers' own state and year of birth.

year; hence γ, λ, δ, and θ are the coefficients on unrestricted age, birth state, birth cohort, and calendar year fixed effects, respectively.[9] The data are collapsed at the state-of-birth by year-of-birth by age level. Observations are weighted by the relevant population in a given cell and standard errors are clustered at the state-of-birth by year-of birth level to account for cohort-specific serial correlation in labor market outcomes.

Given the included fixed effects, the coefficient vector β_a captures deviations from the typical mortality age profiles related to cohort-state-specific variation in the unemployment rate at labor market entry that are uncorrelated with contemporaneous nationwide shocks and independent of responses in graduation timing and migration to state-cohort-specific shocks. Notice that we do not include the current state unemployment rate in our main results; therefore β_a captures the effect of graduating in a recession, given the regular subsequent evolution of local labor market conditions. However, controlling for the current state unemployment rate does not affect results substantially.

3.3 Mincerian specification

We compare our baseline estimates to a specification that links individuals to the economic conditions in the state and the year in which they graduated. Such a specification is often used in previous studies that focused on college graduates and typically use longitudinal data containing direct information on the state and the year of labor market entry. In the CPS, we can proxy for this information using individuals' current state of residence and their reported educational attainment. We construct the "Mincerian" year of labor market entry as year of birth, plus six, plus years of education, and use it as the main explanatory variable in the following regression:

$$outcome_{s,g,t} = \alpha + \beta_e u_{s,g} + \gamma_e + \lambda_s + \delta_g + \theta_t + \pi_d + \epsilon_{s,g,t} \tag{4}$$

where $u_{s,g}$ is the unemployment rate in the state of current residence s at the "Mincerian" year of graduation g. e refers to years of potential experience (years since graduation) and t to the calendar year. We additionally include education-group fixed effects π_d. The data are collapsed by the state of residence, graduation year, calendar year, and four education groups, and observations are weighted by the population weighted cell-size. Standard errors are clustered at the level of graduation year by state.

[9] We drop one dummy for each age, birth state, birth cohort, and calendar year effects. It is well known that cohort, year, and age effects are not separately identified, and hence we drop a second dummy from the birth cohort effects. Choosing alternative restrictions has no bearing on our main results.

Compared to the double-weighted average unemployment rate, the unemployment rate used in the Mincerian specification contains more identifying variation as it is not averaged across multiple years and all states. Moreover, the sample of the Mincerian specification is larger, as the double-weighted specification links each cohort to unemployment rates between ages 16 and 22. While the contrast with the Mincerian model is useful when we analyze economic outcomes, as discussed above the double-weighted specification is our preferred model for two key reason. It is not prone to bias stemming from endogenous timing or migration before graduation or to migration after graduation. Most importantly, the double-weighted specification can be used in the Vital Statistics mortality data.

3.4 Sample restrictions

State-level unemployment rates are available from the Bureau of Labor Statistics only since 1976. Therefore, we exclude individuals who were born before 1960, i.e., who were 16-year-olds before 1976, in the double-weighted specification and individuals who graduated before 1976 when using the Mincerian specification. Our sample period starts in 1979, which is the year in which Vital Statistics death certificates started to report the state of birth. All data sets are further restricted to individuals born in the United States, with the exception of CPS data prior to 1994. Finally, we include only cohorts that are observed across all ages, resulting in different cohort ranges when analyzing wider or narrower age ranges. Our baseline cohorts were born between 1960 and 1967, and are observed from age 19 to age 49.

3.5 Summary statistics

Table 1 shows mortality rates across five-year age groups and mortality causes for the 1960–1967 birth cohorts. Between age groups 19–25 and 46–49, overall mortality increases more than threefold (from 11.3 to 37 deaths per 10,000). Disease-related mortality, which increases from 2.7 deaths to 28.7 deaths per 10,000, drives this increase. Deaths due to external causes fluctuate between 8.6 and 6.1 per 10,000 over these age ranges (see Appendix Figure A.1 for the typical age profiles of mortality due to diseases and external causes). The final column of Table 1 shows mortality due to deaths of despair, which include deaths due to liver disease, suicide, and drug poisoning. Deaths of despair increase from 1.8 to 6.9 per 10,000 over this age range. Part of this increase is driven by the surge of the opioid epidemic and drug poisoning deaths in later years of the sample rather than reflecting age effects. Table A.2 reports summary statistics for labor market and family formation outcomes and displays familiar life-cycle gradients. Tables A.1 and A.2 in the Appendix report summary statistics

by gender and race for mortality and socioeconomic outcomes, respectively.

4 Impacts on mortality

As backdrop to our regression analysis, we assessed nationwide cohort-level trends of overall mortality and mortality due to deaths of despair. Figure 1 shows age profiles for different cohorts that are divided into boom, bust, and regular cohorts depending on the unemployment rate around their graduation age (proxied by age 18 in this figure). As benchmark, dashed lines connect mortality levels at age 40 of boom cohorts. For overall mortality, the dashed line has a negative slope, suggesting that all-cause mortality at age 40 fell over time. For deaths of despair, however, the dashed line has a positive slope, indicating that mortality for those causes increased across cohorts. What is key for our analysis, the mortality rate at age 40 for bust cohorts lies above the dashed line in both figures, indicating that cohorts facing higher unemployment rates at age 18 experienced higher mortality rates at age 40 than those with more favorable economic conditions. At the same time, these figures also show the presence of cohort and time effects unrelated to labor market entry conditions, such as the HIV/AIDS epidemic in the mid-1990s. This highlights the need to control for time and cohort effects in a regression framework if one seeks to estimate the causal effect of economic conditions at labor market entry on subsequent mortality.

4.1 Overall mortality

Figure 2 plots the coefficients (with 95 percent confidence intervals) from regressions of overall deaths per 10,000 on the double-weighted average graduation year unemployment rate , interacted with two-year age groups (β_a in Equation (3)). The effects at age 19 and 20, i.e. the contemporaneous effects of higher unemployment on mortality, are negative. For every additional percentage point in unemployment at graduation, mortality at age 19 and 20 is reduced by about 0.2 deaths per 10,000. This finding is in line with Ruhm (2000) and related studies, who have shown that mortality tends to decrease during times of high unemployment. The figure shows that this negative effect vanishes at subsequent ages, and fluctuates around zero between the mid-20s and the mid-30s. In the late 30s, however, a positive effect on mortality appears that further increases up to the late 40s (see Appendix Figure A.2 for log-mortality results). Mortality at age 49 increases by about 1 death per 10,000 for every percentage point increase in the graduation year unemployment rate.

Table 2 shows the point estimates corresponding to Figure 2 (column 5). To obtain

a balanced sample across all ages our main results focus on individuals born between 1960 and 1967 that entered the labor market approximately between 1978 and 1989. However, this cohort restriction does not drive our main finding. The remaining columns of Table 2 show findings from an increasingly larger number of cohorts, including only the age effects that can be identified with all cohorts present (see also Figure (A.4)). Effects are similar when we shorten the analyzed age window in order to include additional birth cohorts. These results suggest that entering the labor market during bad economic times has a positive long-term effect on mortality. Furthermore, this effect is not visible or even reversed in the short run.[10]

4.2 Effects on cause-specific mortality

Figure 3 shows separate regression results for deaths due to diseases and external causes. It is apparent that the negative effect on mortality at young ages is entirely driven by external causes, while the positive and increasing effects at middle ages is driven by disease-related deaths.

In order to further explore the drivers of the positive long-term effects at middle ages, Table 3 shows regression results for leading disease-related mortality causes. These regressions use mortality in logs as rates differ up to a factor 100 across age groups (see Table 1). For deaths due to liver disease, heart disease, and lung cancer we find an effect pattern that resembles the positive effects for overall disease-related mortality in Figure 3. We do not find effects of early labor market conditions on deaths due to other cancers, even though cancers other than lung cancer are jointly a leading cause of death in the late 40s.

Overall, we find that the positive effects on mortality in middle age from adverse initial labor market conditions are to a large extent driven by disease-related causes that can be linked to detrimental health behaviors, such as smoking, drinking, poor diet, and lack of exercise.

Yet, it turns out that the apparent absence of an effect on overall external deaths in the right panel of Figure 3 masks important heterogeneity across different causes. Figure 4 zooms in on different external causes of death. The negative contemporaneous effects at ages 19 and 20 are entirely driven by a decline in accidents (first panel).[11] The second panel of Figure 3 shows that there is no effect on suicide, either at younger or at middle ages. For drug poisoning deaths (third panel), however, a positive and significant effect appears after

[10] Our findings are robust to the choice of cohort, age, and year fixed effects. Appendix Figure A.3 shows that specifications with alternative cohort effects result in virtually the same effect pattern.

[11] This finding is in line with Evans and Moore (2011, 2012) who show that economic activity increases short-term mortality.

age 40.

Figure 5 shows positive midlife impacts on deaths of despair, as defined by Case and Deaton (2015, 2017), and compares effects to broader sets of deaths related to health behaviors, including lung cancer and heart disease. The left panel shows effects on deaths per 10,000. Adding lung cancer has little impact on estimates at younger ages but greater impacts in the 40s. At age 49, the total impact doubles compared to deaths of despair alone. Further adding heart disease increases impacts more dramatically, with point estimates around 0.6 in the late 40s, about six times the impact on deaths of despair alone. Jointly, these causes of death can explain approximately two thirds of the overall increase in mortality we find (see Figure 2). Heart disease is the leading cause of death in this age group, which may partly explain the large impacts when this cause is included. The right panel of Figure 5 therefore shows effects on log mortality rates, which can be interpreted as percent increases. The resulting differences across the three groups of causes are smaller, but the specification that includes heart disease still yields the largest point estimates.

Overall, we interpret the patterns shown in Figure 5 to reinforce our finding from Table 3 that the positive effects on mortality in middle age from adverse initial labor market conditions are to a large extent driven by disease-related causes that can be linked to detrimental health behaviors. The results also indicate that mortality analyses that focus on long-term socioeconomic decline and the role of economic hardship might include lung cancer and heart disease deaths, besides the more strictly defined deaths of despair.

4.3 Demographic heterogeneity of mortality effects

Table 4 shows impacts of the unemployment rate at the time of labor market entry on mortality by demographic group (summary statistics reported in Table A.1). Given sizable differences in mortality levels across race and gender, we show effects on log mortality.

Midlife impacts of the unemployment rate at labor market entry on overall mortality are similar across race, but they are two to three times larger for women than for men in percentage terms (albeit relative to a smaller baseline rate). Interestingly, for non-whites, some mortality impacts already appear in the 20s, and this early effect occurs both for disease-related deaths and deaths of despair. This could be partly driven by larger adverse consequences of initial unemployment rates on earnings and poverty for non-whites (Schwandt and von Wachter (2019)). In addition, both the HIV and the crack epidemics of the early 1980s to mid-1990s hit the non-white community much harder. During that time, HIV-infections and crack addiction could lead to death within a few years. Appendix Table A.5 shows disease-

related mortality excluding HIV/Aids deaths, and indeed, point estimates for non-whites are reduced by about one-third.

Midlife impacts on deaths of despair in Panel (B) of Table 4, on the other hand, are driven by whites. Point estimates for non-whites are small and not significantly different from zero. This finding is in line with literature showing that non-whites were less strongly hit by the opioid epidemic as doctors discriminated against non-white patients in their prescribing behavior (Pletcher et al., 2008). When we include lung cancer and heart disease in the second panel, we find, if anything, stronger midlife impacts for nonwhites.

In sum, the pattern in Table 4 suggests that while the overall mortality impact of initial labor market conditions is similar across races, there are differences in the most-affected causes, partly depending on the exposure of different demographic groups to different epidemics occurring during the sample period.

4.4 Magnitude of mortality effects

How large are the estimated mortality effects? The answer to this question crucially depends on whether and how we assume the mortality impact sustains after age 49, the end of our age window. If mortality impacts only occurred between ages 19 and 49, the impact on life expectancy would be moderate: entering the labor market during the 1982 recession (i.e., facing a 3.9 percentage point increase in the entry unemployment rate) would reduce life expectancy by 0.6 months. However, given our estimates of increasing mortality impacts over age, it seems more likely that effects sustain past age 49. We calculate life expectancy losses based on a linear and a percent extrapolation (illustrated in Figure A.5). The linear extrapolation assumes that impacts increase linearly with age, following the trend estimated for age 40 to 49. For the percent extrapolation we use a constant 2 percent effect for the rest of life.[12]

Adding the impact of a 3.9 percentage point increase in the entry unemployment rate to the mortality age profile (using the percent extrapolation) makes only a minor difference at any given age (see Figure A.6). But these impacts accumulate over the life cycle. As reported in Table 5, the linear and percent extrapolations imply a loss in life expectancy of 5.9 and 8.9 months, respectively. For a hypothetical cohort of 4 million (approximately the

[12] Alternatively, we can extrapolate impacts at older ages based on log-linear Gompertz models. However, such models imply much larger impacts at older ages than the linear or percent extrapolations that we use. Fitting a linear age trend through the log mortality impact (Figure A.2, right panel), with a negative impact at age 19 and positive impacts above age 40, results in a positive age slope implying that the impacts on log mortality increase with age, beyond the 2 percent observed at the late 40s.

size of the 1964 birth cohort), this is equivalent to a total of 2 million and 3 million life years lost, respectively.

A life expectancy loss of six to nine months is reasonable compared to mortality impacts of labor market and business cycle shocks estimated in previous literature. Sullivan and von Wachter (2009) find that job loss as part of a mass layoff reduces life expectancy by 1 to 1.5 years. Our mortality effect is about half the size, but our estimated impacts on income are smaller, too. Sullivan and von Wachter (2009) estimate a 10–20 percent reduction in lifetime earnings, whereas results in Schwandt and von Wachter (2019) imply a loss of 6% of cumulated earnings in the first ten working years from entering the labor market in a recession. Given our findings in Section 5 this is an underestimate of the total effect on earnings.

Another benchmark is the contemporaneous mortality effect of business cycle fluctuations (Ruhm, 2000; Miller et al., 2009; Stevens et al., 2015). We use the contemporaneous impacts by single year of age reported in Miller et al. (2009) to calculate the effect of a 3.9 percentage point increase in the unemployment rate on life expectancy (this refers to the impact on the length of life for an individual that experiences a strong local recession at every year of age throughout life), resulting in an increase in life expectancy by 1.9 months. In other words, the negative longevity impacts of entering the labor market during a recession is three to 4.5 times larger than the positive contemporaneous mortality impacts.

Another way to benchmark our estimates is to compare them the overall increases in deaths of despair studied in the literature. Between 1990 and 2010, deaths of despair increased by around 3.4 per 10,000 at age 50 to 54 (Case and Deaton, 2017). These cohorts entered the labor market in the late 1960s to the late 1970s, with recessions occurring during a total of two years and three months of that period. Even in the hypothetical case that all of these cohorts were fully impacted by a strong recession (experiencing a 4 percentage point increase in the entry unemployment rate), deaths of despair would only increase by 0.4 per 10,000.

5 Impacts on labor market and family formation outcomes

In this section, we expand our analysis and estimate the longer-term effects in middle age of entering the labor market in a recession on a broad set of labor market and family formation outcomes. This is of interest in its own right since most prior studies have focused on effects in the first ten to fifteen years, partly motivated by the desire to combine a larger number of entry cohorts in their estimates. Here we focus on cohorts coming of age before, during,

and after the early 1980s recession and hence can follow these cohorts into middle age. In addition, these findings are relevant since they give further context for the increases in mortality in middle age we find. In interpreting these findings, it is important to bear in mind that our research design allows us to estimate the causal effect of adverse initial entry on a range of outcomes, but does not allow us to disentangle likely complex interdependencies.

To benchmark our findings to the prior literature, we start by comparing the short- and medium-run effects for log earnings for individuals entering the labor market in 1976–1987 (the equivalent to our 1960–1967 baseline birth cohorts in the double-weighted specification) with results for individuals entering 1976–2015 that are studied in Schwandt and von Wachter (2019). We then study a broader set of socioeconomic outcomes in middle age for the 1976–1987 entry cohorts using the double-weighted specification discussed in Section 3.2.

5.1 Earnings impacts using the Mincerian specification

The top panel in Figure 6 shows the impact of the graduation year unemployment rate on log earnings by years of potential labor market experience based on the Mincerian specification in three different samples. The blue and green lines show effects in CPS and Census/ACS data, respectively, for individuals graduating in 1976–1987 who can be followed up to 30 years past graduation. The dashed lines show 95 percent confidence intervals. The red line with hollow squares (labeled SvW) replicates the baseline estimates from Schwandt and von Wachter (2019) using CPS data and including observations from all individuals graduating between 1976 and 2015.[13]

All three sample specifications show a similar short- and medium-term effect. Initially, earnings are decreased by -2.5 to -4 percent for every percentage point increase in the graduation year unemployment rate, and this effect fades out within ten years. But this is not yet the full story. Negative effects reappear 15 to 20 years after graduation and they remain at around minus one percent until the end of our observation period.

As described in Section 3, to address selection issues we also replicated these findings using our double-weighted unemployment rate. The bottom panel of Figure 6 shows earnings effects by age using the double-weighted specification and the same baseline cohorts as in the mortality analysis. Compared to the Mincerian specification, estimates are less precise, both due to the smaller number of included cohorts and the reduction in identifying variation

[13] As described in the data section, we use state of current residence in the CPS data as a proxy for state of birth. The similarity between the CPS and the Census results suggests that this proxy is reasonable.

caused by the two-way averaging inherent in the double-weighted average graduation year unemployment rate. Moreover, the effect at age 19 is smaller than the effect at age 20 as at that age those cohort members with more than 12 years of education have not yet entered the labor market. But besides these differences, the overall pattern is very similar to the Mincerian specification shown in the top panel. The initial negative effects of the unemployment rate at labor market entry fade out in the late 20s and early 30s but then reappear in the mid-30s. The point estimates, shown by decade of age in Table 6 (A), imply about a one percent loss in middle-age income for every percentage point increase in the double-weighted graduation year unemployment rate.

5.2 Effects on labor market outcomes

For the analysis of additional outcomes, we pool the CPS and Census/ACS samples to maximize our sample sizes. Figure 7 and Table 6 show the impact on log earnings in the pooled sample, along with further labor market outcomes. In midlife, despite having lower annual earnings, recession graduates have higher rates of labor force participation. However, compared to the means in the same age group shown in the bottom panel of Table 1 (78.2 percent), the point estimates shown in Table 6 imply that the effect of a three-point rise in the unemployment rate increases the longer-term employment rate only by 0.92 percent relative to the mean. Recession cohorts are also less likely to receive welfare income in middle age. Yet, despite the decline in earnings and welfare receipt, they do not experience a longer-term impact on poverty rates. We will discuss these results further when considering separate demographic groups.

5.3 Effects on family formation

Figure 8 and Table 6 show that we find precisely estimated effects of entering the labor market in a recession on family formation outcomes over the life course and up to middle age. In the short-term, recession graduates experience slightly elevated marriage rates. However, with increasing age this marriage bonus decreases and turns negative in the late 30s when recession graduates are less likely to be married than their more lucky counterparts. The second panel in Figure 8 shows that the effect on marriage rates is partially driven by divorces, which spike in the late 20s and follow the peak of the marriage rate effect by about five years. For unlucky labor market entrants, divorce rates remain elevated until age 49. Compared to the means at ages 40–49 (bottom panel of Table 1), the magnitude of these effects (reported in Table 6) is small to moderate. A three-point rise in the unemployment rate — roughly

the typical increase in a recession — would imply a reduction of somewhat less than 2.5 percent in middle age marriage rates relative to the mean (63.5 percent), and a rise of about 3.5 percent in middle-age divorce rates relative to the mean (19.34 percent).

The bottom two panels in Figure 8 show corresponding effects of entering the labor market in a recession on cohabitation with own children. Recession graduates are more likely to have own children living in their household in their mid-20s, in line with higher marriage rates at that age.[14] However, in midlife they are less likely to have any own children in their household. The effects are again of moderate size, with a three-point rise in unemployment rates implying effects of 3-4 percent relative to the mean. We cannot distinguish whether this effect is driven by adolescent children moving out, younger children living with a divorced partner, or lower overall fertility rates. Given that Currie and Schwandt (2014) show increased unemployment rates during women's early 20s lower their completed fertility, it is likely that at least part of our findings is due to reduced overall fertility rather than being only a reflection of differences in birth spacing or cohabitation patterns.

5.4 Demographic heterogeneity of socioeconomic impacts

In Tables 7 and 8, we analyze the impacts on labor market and family formation outcomes separately by gender and again two broad race groups, whites (referring to non-Hispanic whites), and non-whites (including Hispanics). The main takeaway is that even though the initial labor market effects are much stronger for non-whites, the long-term effects on income and family formation tend to be driven by whites. While both white men and women experience a reduction in midlife marriage rates and in the presence of children in household, longer-term effects on earnings are concentrated on white men. While key findings are precisely estimated, an important caveat is that in some instances we do not have sufficient precision in the non-white samples.

The initial earnings effect of entering the labor market in a recession is about double for non-whites compared to whites (Panel A of Table 7). Both groups experience a reduction in the initial loss, and by age 30–40, non-whites have no long-term loss. However, the losses for white men never quite disappear in their thirties and the losses grow larger in their 40s. To the extent that negative long-term effects for whites may be driven by employment in jobs

[14]Note that the zero effect at age 19, one year after the modal labor market entry, is somewhat at odds with the well established finding of procyclical fertility (Currie and Schwandt, 2014). One potential explanation is that fertility tends to lead rather than lag fluctuations in the unemployment rate (Buckles et al., 2018). It is also likely that upon exiting school around age 18, individuals take some time to learn about the adverse labor market situation and to find a partner and have children, leading to a gradual increase of the effect by age.

with flatter income age profiles (see Section 6), the absence of an effect for non-whites could mean that jobs with steep income profiles are not available to disadvantaged populations even in better economic times. Among these cohorts, white women appear to have smaller, imprecisely estimated initial earnings losses and no signs of longer-term effects.

In the bottom row of Table 7, we show the p-values of two-sided t-tests of the equality of the coefficients at age 40-50 of each demographic with the corresponding effect for white men. It is clear that white men have statistically different long-term earnings effects from entering the labor market than women and non-white men. Women, however, appear to experience stronger effects on labor supply (Panel A of Table 7).[15] As a mirror image, welfare receipt is significantly negatively impacted for all demographic groups except for white men (Panel B of Table 7). The effects are strongest for non-white women, who also have substantially higher baseline rates (Table A.2). There are no precisely estimated effects on poverty for any demographic group. The findings on labor supply, welfare receipt, and poverty could be partly influenced by the strong 1990s labor market that affected these cohorts in their mid- to late-30s and early 40s. As further discussed in Section 6, for women both the short- and longer-run labor supply effects could also be affected by family considerations.

Turning to family status, Panel A of Table 8 shows that the effect of entering the labor market in a recession on marriage and divorce rates are largely concentrated on white men and women both in the short and long run. Marriage rates are strongly elevated among white recession graduates in their 20s, perhaps pointing to lower opportunity costs of marriage and family formation. However, this marriage premium is lost quickly due to higher divorce rates in the late 20s (for men) or the catch up in marriage rates of non-recession cohorts. In midlife, marriage rates are consistently lower among white recession graduates. White men in particular also face permanently higher divorce rates.[16]

Finally, cohabitation with any own children and the number of cohabiting children is impacted negatively in midlife for both white men and women (Panel B of Table 8).[17] Interestingly, however, female recession graduates are initially more likely to report own children

[15] White women, and non-white women in particular, experience large reductions in initial labor supply upon entering the labor market in a recession. Interestingly, especially for white women, these effects turn positive in their 30s and they remain higher in their 40s, potentially linked to lower marriage rates across those ages (Johnson and Skinner, 1986). The effects are of the same magnitude but imprecisely estimated for non-white women. Non-white men also experience positive midlife labor supply effects.

[16] The average differences in the effects on marriage rates in the short and long run between whites and non-whites are precisely estimated. The p-value of a two-sided t-test of equality of the effect for the two race groups is 0.04. In contrast, the corresponding p-value for a test of the effect on any children living in the household is 0.15.

[17] While non-whites experience imprecisely estimated negative effect on the number of cohabiting children in the long run, there is little effect on likelihood of living with any children.

in their household, and this positive impact lasts until their mid-30s. As it is unlikely that non-recession female graduates entirely catch-up in their late 30s when female fertility is low, this pattern suggests that part of the negative long-term effect on cohabitation with own children is driven by earlier-born offspring leaving the household. The pattern would be consistent with a rise in early life marriage rates observed at least for white women. However, here we cannot conclusively say how these effects relate to changes in spacing and total fertility.

6 Discussion

Our analysis has revealed a range of short- and long-term impacts of labor market entry conditions on mortality and socioeconomic outcomes. In this section, we discuss potential underlying mechanisms. While identifying the effects of specific channels goes beyond the scope of this paper, the discussion relates our results to economic hypotheses and findings in the literature.

6.1 Discussion of mortality effects

The positive short-term effects of entering the labor market in a recession on longevity we find are in line with an established literature that has documented that mortality is procyclical, i.e., the contemporaneous effects of aggregate unemployment on mortality are negative (Ruhm, 2000; Miller et al., 2009; Stevens et al., 2015; Ruhm, 2015). For labor market entrants, the group of primary concern in this paper, we find these contemporaneous effects are relatively short-lived and fade within a few years after labor market entry.[18]

While we find no effects in the 20s and early 30s from entering the labor market in a recession, positive mortality impacts from initial unemployment rates appear in the late 30s and grow steadily in their 40s, and up to age 49. These results complement previous studies that find long-run mortality effects for economic and labor market conditions around birth, adolescence, and before retirement (Van den Berg et al., 2006; Cutler et al., 2016; Coile et al., 2014). Our results are also consistent with findings that empirical SES-health gradients increase with age up to age 55 (e.g., Case et al., 2002; Currie and Stabile, 2003). While it has been difficult to establish how much of the gradient and its increase with age are due to causal effects, our results show that labor market entry shocks do have such increasing

[18]For these individuals, we show the contemporaneous mortality effects are mainly driven by accidents, suggesting that depressed economic activity lowers the risk of fatal injuries on the road and at the workplace (in line with Evans and Moore (2011, 2012)).

causal effects on mortality. The patterns we find are predicted by the model of life-cycle mortality by Moreau and Lleras-Muney (2019) and can be generated both by temporary and permanent reductions in health investments. In their model, even temporary reductions in health investments have permanent scarring effects because health depreciation over time is modeled as additive, not multiplicative in the stock. Notably, in either case the model predicts low impacts in young adulthood when accumulated age-driven appreciation is still low, pushing few individuals over the death threshold.

Our finding that these increases are driven by causes linked to health behaviors, suggests that entry conditions affect midlife mortality via impacts on SES and living circumstances, which in turn affect health (rather than a direct, non-health effect on mortality via, for instance, a rise in accidents). One mechanism driving these long-term mortality increases could be persistent behavioral or psychological responses to the initial strong impacts on economic outcomes. At an especially impressionable age and a vulnerable transition period, a person may be more likely to adopt unhealthy behaviors or fail to eliminate those acquired in high school or college (Kessler et al., 2005).

The persistent earnings reductions and worse social outcomes we find in the long run may further contribute to poor health behaviors and unhealthy living circumstances, e.g., via social isolation (Cornwell and Waite, 2009). These factors might drag down health sufficiently to raise mortality in midlife (Steptoe et al., 2013), when the incidence of adverse health generally increases. As further discussed below, poor health might in turn reinforce socioeconomic impacts, lowering an individual's labor market productivity and socio-economic status and worsening social isolation.

Our results provide empirical support for the hypothesis that increases in deaths of despair can be driven by persistent effects of socioeconomic decline (Case and Deaton, 2015, 2017). In particular, Case and Deaton (2018) emphasize the role of a broad set of socioeconomic outcomes, including family formation and cohabitation, and conjecture that the full impact of some of these factors may only occur over a longer time horizon. Our results show this is the case for unlucky labor market entrants. While our analysis is able to identify drivers of deaths of despair linked to market entrants' exposure to local short-term business cycle fluctuations, Pierce and Schott (forthcoming) and Autor et al. (forthcoming) document impacts of longer-term structural changes, such as reductions in manufacturing jobs induced by globalization and import competition.

In a discussion of Case and Deaton (2017), Lleras-Muney (2017) examines the trends in deaths of despair in light of a structural framework of life cycle mortality (based on Moreau

and Lleras-Muney, 2019). Lleras-Muney (2017) concludes that midlife increases in deaths of despair "can only be the result of permanent changes in parameters beginning early in adulthood" and that "there [must have] been a permanent deterioration in one or more factors that affect health, starting at about the time of labor market entry." (As noted above, a permanent health deterioration can also be generated a temporary reduction in investments.) Our results demonstrate that labor market entry conditions are one such factor with persistent impacts on midlife mortality.

At a broader level, our results also contribute to an ongoing, highly publicized debate about mortality trends and inequality in mortality. Overall, mortality has strongly declined over the past decades, with particular improvements for African-Americans (Currie and Schwandt, 2016b). In contrast, poor, non-hispanic whites in middle age have faced stagnating or even raising mortality rates (Case and Deaton, 2015). Moreover, Currie and Schwandt (2016a,b) document strong convergence in mortality rates between rich and poor areas for infants, children, and adolescents, while inequality in mortality has stagnated or increased at middle and older ages (Wilmoth et al., 2011; Pijoan-Mas and Ríos-Rull, 2014; Goldring et al., 2016; Chetty et al., 2016). Our results suggest that initial labor market conditions might play a role in transmitting the effects of economic conditions on inequality in mortality.

6.2 Discussion of socioeconomic effects

Several hypotheses have been advanced in the literature to explain the short- to medium-term effects on earnings from entering the labor market in a recession. Short-term earnings effects have been explained by a combination of worsening job opportunities, time-intensive job search, and reduced opportunities for skill investment. For example, Oreopoulos et al. (2012) suggest persistent effects can arise if time-intensive job search slows down as workers mobility costs rise with age.[19] To explain *worsening* effects on earnings of early labor market conditions that initially seemed to fade documented by Kahn (2010) for college graduates

[19]Oreopoulos et al. (2012) lay out a theoretical model in which high- and low-skill workers search for jobs, and each aim to obtain jobs at higher paying firms. A temporary reduction in the availability of good jobs leads to a loss in earnings as workers are temporarily downgraded to worse firms. While typical measures of search intensity would suggest a catch-up within three to four years once the economy returns to normal, the pattern can take longer if workers search costs increases as they age (for example due to family formation). The patterns could last longer if workers have the opportunity to accumulate firm- or occupation-specific skills once they found the right employer. If better firms also offer better opportunities for career advancement or human capital accumulation, then initial effects can last a long time. Typical models of career development have concave experience or tenure profiles in wages, and hence do not predict *permanent* effects on earnings from adverse initial conditions. To obtain such permanent effects, Oreopoulos et al. (2012) posit that recovery may be incomplete if workers stop searching or stop investing in skills, perhaps because the costs of doing so rise with age, e.g., due to family formation or homeownership.

and we confirm for all white men, Gibbons and Waldman (2006) propose a model in which firms are ex-ante identical, but decide to offer certain jobs with more opportunities for career advancement only in good economic times. As a result, unlucky entrants could remain stuck in lower quality jobs, from which they may be more likely laid off again. Similarly, if one were to introduce layoffs in the Oreopoulos et al. (2012) search environment, otherwise equal unlucky cohorts have, on average, lower job tenure, and hence may again be most at risk of experiencing a layoff in middle age. Given that workers graduating in the 1982 recession faced the early 2000s recession in their 40s, this may be a plausible explanation. The fact that we do not observe employment effects is compatible with this explanation, since most job losers find jobs again quite quickly.

Another reason for a widening gap in earnings may be adverse health shocks themselves that typically start to occur in middle age. For example, we find that the effect of entering the labor market in a recession on mortality starts to rise in the mid-30s (e.g., see disease-related deaths in Figure 2). If these increases in mortality reflect the overall health of these cohorts, such lower health would likely reduce overall productivity and hence earnings (Meyer and Mok, 2019). Worsening health may also affect family stability and, to a lesser extent given the age range, fertility (Balen and Crawshaw, 2006). It is important to note that any worsening health does not appear to lead to reductions in employment, as we find affected workers' earnings rise in their 30s and 40s, on average, compared to luckier labor market entrants.

It is likely that at least part of the changes in broader socioeconomic outcomes we studied are themselves directly affected by initial labor market conditions. On the one hand, it is possible that workers respond to the initial shock in ways that have repercussions over the life cycle. For example, in the short run, individuals may decide to delay marriage and fertility due to lower income. In contrast, a lower opportunity cost, especially for women, may lead to anticipated marriage, consistent with our findings. In the long run, these marginal marriages may provide less stable matches or may be more unstable due to lower lifetime earnings and wealth. On the other hand, the reductions in earnings we find may also directly affect workers' opportunities in the marriage market, and hence their choices to get married and have children.[20]

[20]Some of our group-specific findings could also be explained by economic mechanisms. For example, while for women a recession-induced reduction of the opportunity cost of childbearing may initially lead to increased family formation, in the longer run, the negative impacts on marriage and cohabitation rates with own children could reduce the opportunity cost of employment, and thus drive female recession graduates into the labor market at older ages. Another hypothesis could be that non-whites have lower marriage and higher out-of-wedlock birth rates, as well as no long-term income effects (due to lack of steep career ladders), which is why family formation outcomes are not impacted.

Given that socioeconomic outcomes and earnings are likely to be to some degree directly affected by early initial conditions, they then in turn might themselves affect health and hence mortality. For example, both lower lifetime earnings and wealth, as well as higher divorce rates and smaller families, may worsen health and increase mortality. While this is an interesting question for future research, here we cannot tease apart the likely complex causal pathways running from health to socioeconomic status and vice versa. However, our finding clearly indicate that early labor market entry has a broad effect on individuals lives up until age 50, affecting both mortality as well as family formation and labor market outcomes.

7 Conclusion

In this paper we have used several large cross-sectional data sources and a new approach to estimate midlife effects of entering the labor market in a recession on mortality by cause and various measures of socioeconomic status. This is the first paper to study adverse labor market entry on mortality and on socioeconomic status effects going beyond age forty. To obtain such longer-term effects, we focus on cohorts coming of age in the late 1970s to mid-1980s, a period of particularly pronounced labor market swings. To estimate the unemployment rate in the state of labor market entry based on cross-sectional data despite data limitations and selection from mobility, we implement a new approach based on a double-weighted unemployment rate introduced by Schwandt and von Wachter (2019).

We obtain four main findings. For cohorts coming of age during the early 1980s recession, a temporarily higher state unemployment rate at the age of labor market entry leads to precisely estimated increases in mortality that appear in the late thirties and increase until age 50. These increases in mortality are driven to an important extent by a rise in both disease-related and "external" causes, including lung cancer, liver disease, and drug poisoning. We also find entering the labor market during a recession has a substantial impact on a broad range of measures of socioeconomic status in middle age, including a decline in marriage rates, a rise in divorce rates, and a decline in family size. We also find that after initial recovery in their mid-thirties, adversely affected entry cohorts suffer a reduction in earnings as they reach their mid-forties. Finally, while the effects on overall mortality are similar by race, increases in deaths of despair appear to be chiefly concentrated among white, non-Hispanic men. White men also tend to experience a decline in earnings in midlife and tend to experience larger reductions in family stability than their non-white counterparts. This is despite the fact that non-whites experience larger short-run effects on earnings and other outcomes (Schwandt and von Wachter, 2019).

These findings suggest long-term costs of entering the labor market in a recession may be greater than previously understood in the related economic literature. In particular, our finding that adverse initial labor market conditions raise long-term mortality implies that the costs of recessions go beyond initial earnings effects documented in the literature. They confirm findings from a growing body of literature that labor market shocks more generally can have long-term consequences for health and mortality (e.g., Sullivan and von Wachter, 2009, Cutler et al. (2016)). Our result that various socioeconomic outcomes also worsen in midlife further underscores that initial labor market conditions have effects that can last beyond the initial adjustment period typically studied. Moreover, mortality can be viewed as an extreme health outcome, not capturing all impacts on latent health. Our estimates are therefore likely to provide a lower bound of the full long-term health effects of initial economic conditions for the cohorts we study.

Addressing a broader literature on trends in health inequality and the relationship of socioeconomic status and health, our results provide evidence of a positive causal pathway from economic conditions to health and mortality that becomes relevant as individuals age. We find more than a decade needs to pass before these effects are measured in the data. These results suggest that some of the roots of the trends in health inequality observed in midlife and old-age might be found in earlier periods of life, consistent with evidence in Case et al. (2002) and Currie and Stabile (2003) and predictions of a life cycle models of mortality in Moreau and Lleras-Muney (2019). Mortality can be viewed as an extreme health outcome, not capturing all impacts on latent health. Hence, ur estimates may provide a lower bound of the full long-term health effects of initial economic conditions for the cohorts we study.

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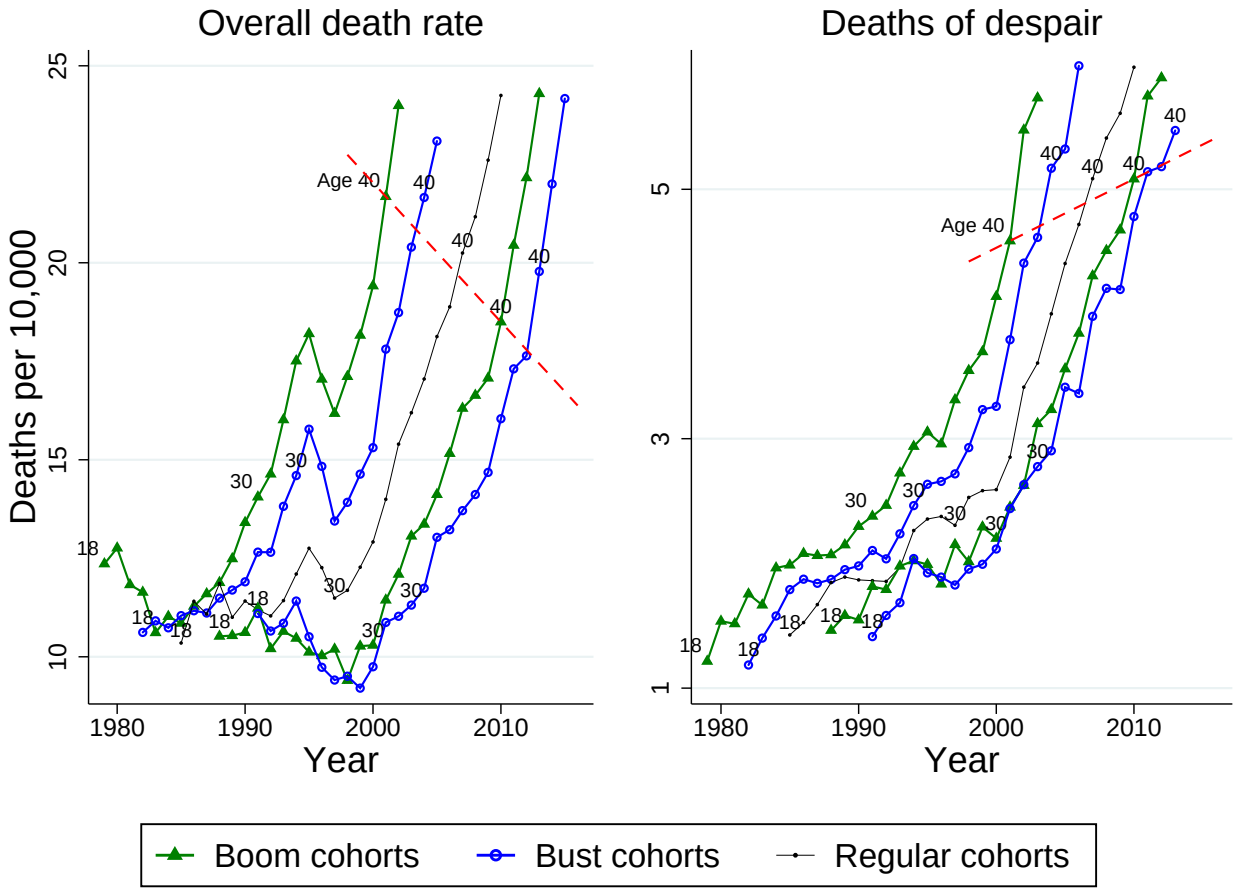
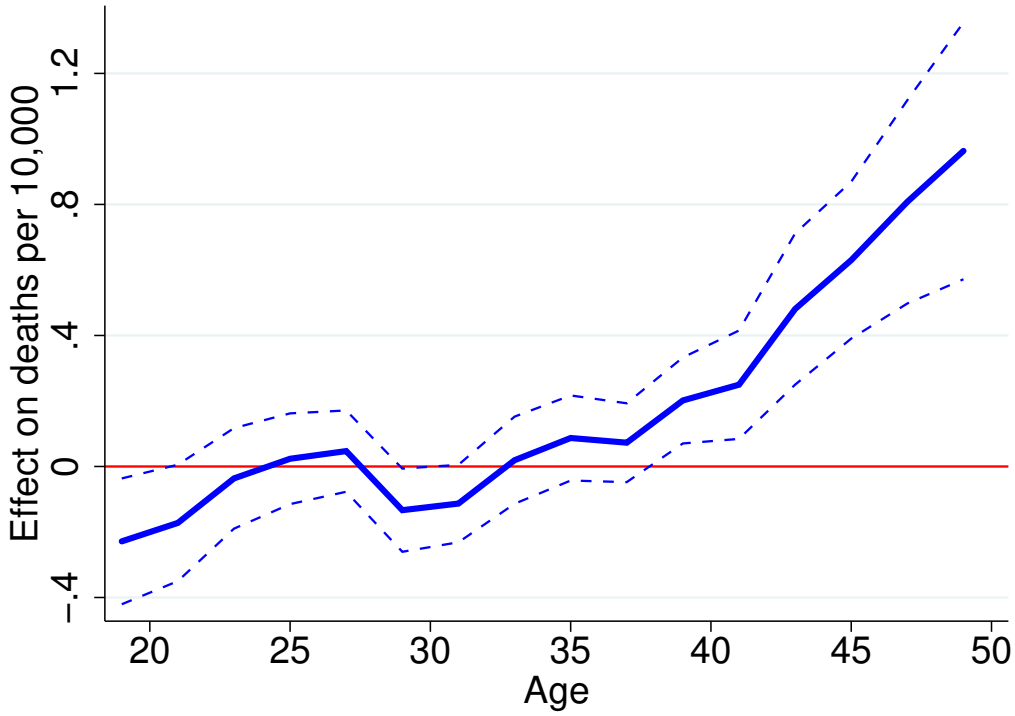
8 Figures and Tables

Figure 1: Mortality Profiles and US-wide Unemployment Rates at Age 18, Across Cohorts



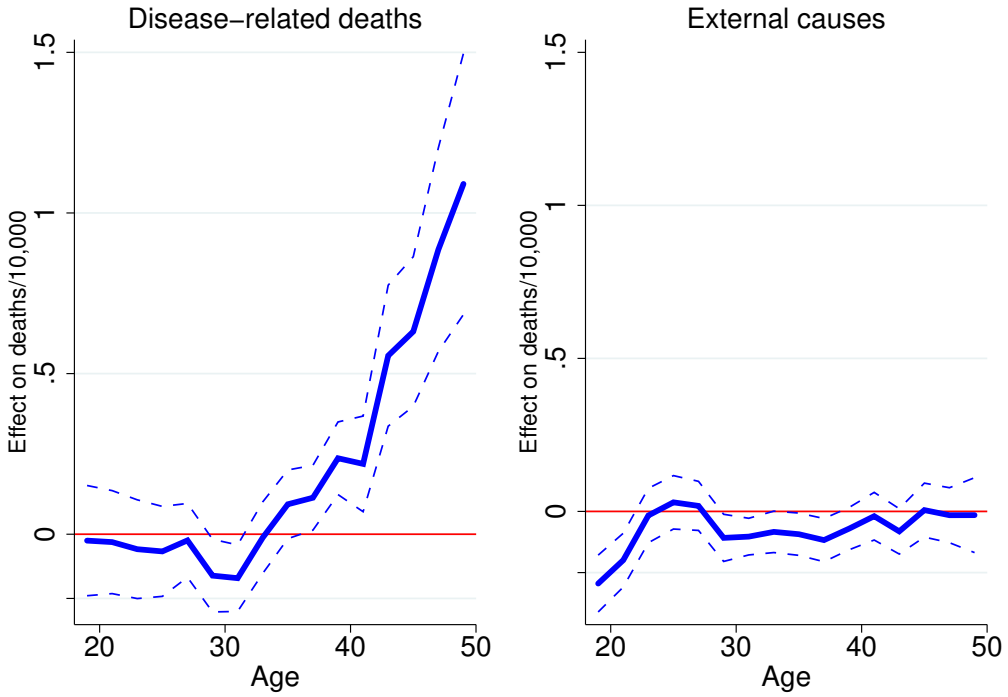
Notes: Mortality profiles from age 18 up to age 43 are shown for selected cohorts. Boom/bust/regular cohorts refer to cohorts experiencing low/high/medium national unemployment rates at age 18. Dashed red lines fitted through the mortality rates of boom cohorts at age 40.

Figure 2: Effect of State Unemployment Rate at Labor Market Entry on Mortality Rate by Age



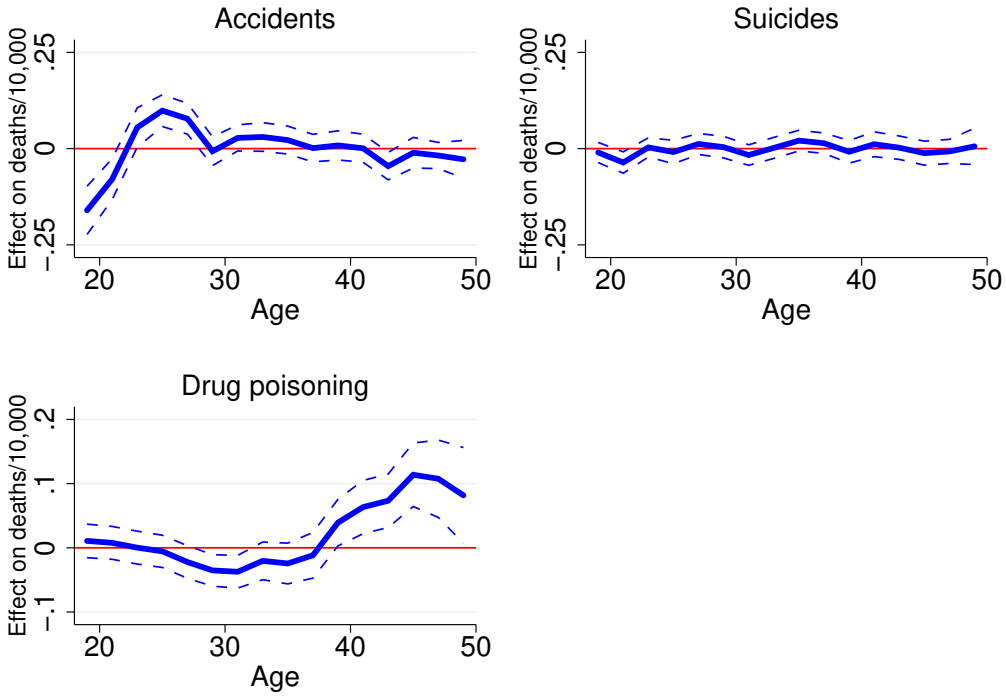
Note: Coefficients and 95 percent confidence intervals are plotted from a regression of death rates on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Vital Stats, Census, and ACS, years 1979–2016. Further regression and sample details as in Table 2. Coefficients for 5-year age groups are reported in Table 2 column (5).

Figure 3: Effect of State Unemployment Rate at Labor Market Entry on Cause-Specific Mortality



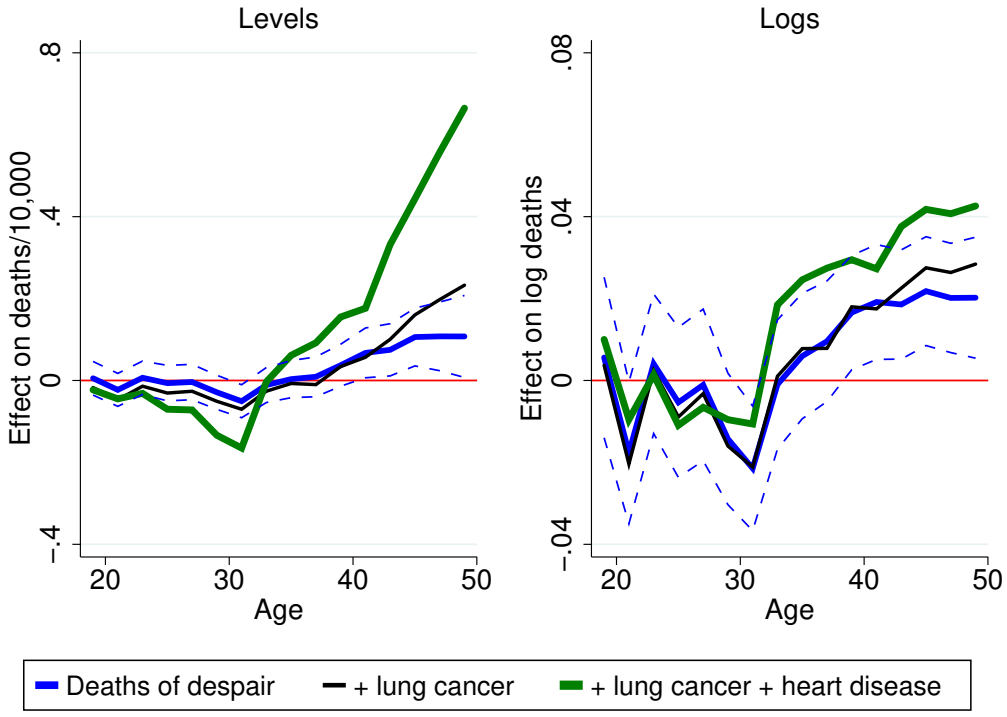
Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of cause-specific death rates on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Vital Stats, Census, and ACS, years 1979–2016. External causes refer to all causes not included in disease-related deaths. Further regression and sample details as in Table 2.

Figure 4: Effect of State Unemployment Rate at Labor Market Entry on external cause mortality and deaths of despair



Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of cause-specific death rates on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Vital Stats, Census, and ACS, years 1979–2016. Accidental drug poisoning counted as "drug poisoning" but not as "accidents." Further regression and sample details as in Table 2.

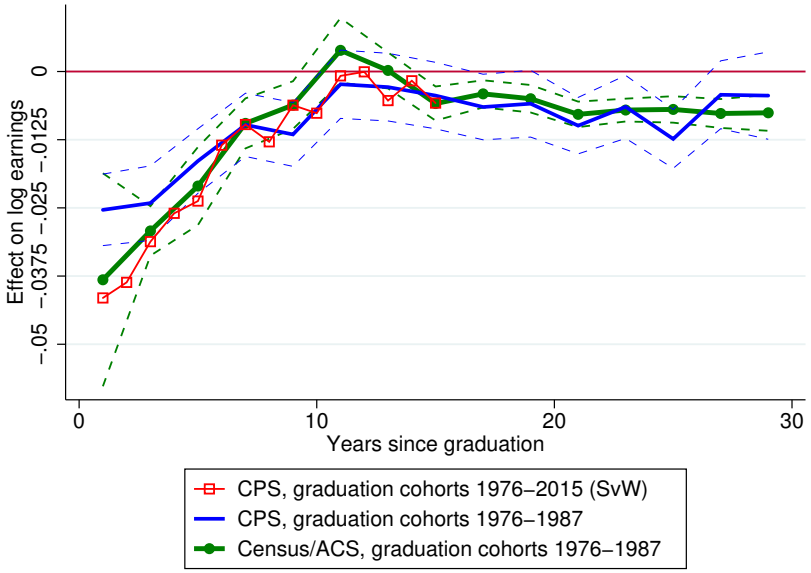
Figure 5: Effect of State Unemployment Rate at Labor Market Entry on "deaths of despair" and related causes



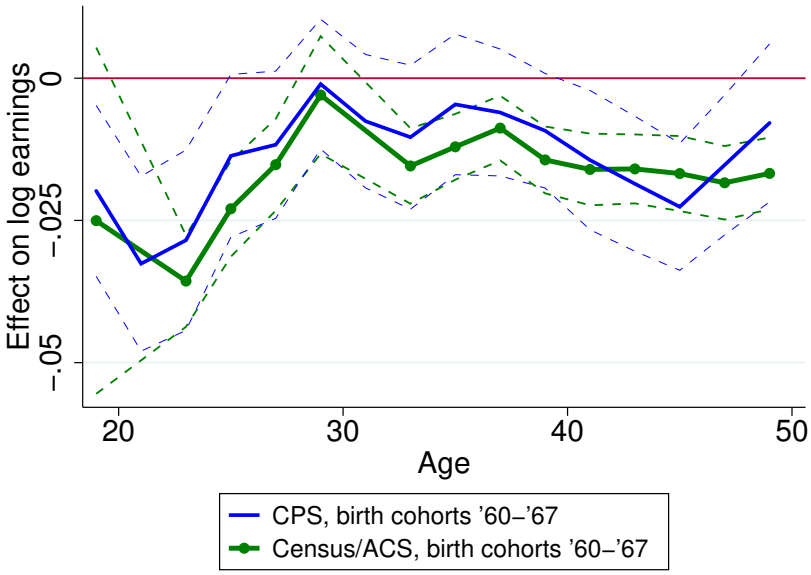
Note: Coefficients and 95 percent confidence intervals are plotted from regressions of cause-specific death rates on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Vital Stats, Census, and ACS, years 1979–2016. Accidental drug poisoning counted as "drug poisoning" but not as "accidents." Further regression and sample details as in Table 2. Deaths of despair include deaths due to liver disease, drug overdoses, and suicide (blue line). The thin black line additionally includes deaths due to lung cancer. The thick green line additionally includes deaths due to lung cancer and heart disease. Because of the sharp increase in deaths due to lung cancer and heart disease with age, we show the effects in both levels and logs. We only show confidence intervals for deaths of despair to avoid clutter.

Figure 6: Effect of State Unemployment Rate at Labor Market Entry on Earnings by Cohort and Data Set

Panel A: Mincerian Specification – Effects by Potential Labor Market Experience



Panel B: Double-Weighted Specification – Effects by Age

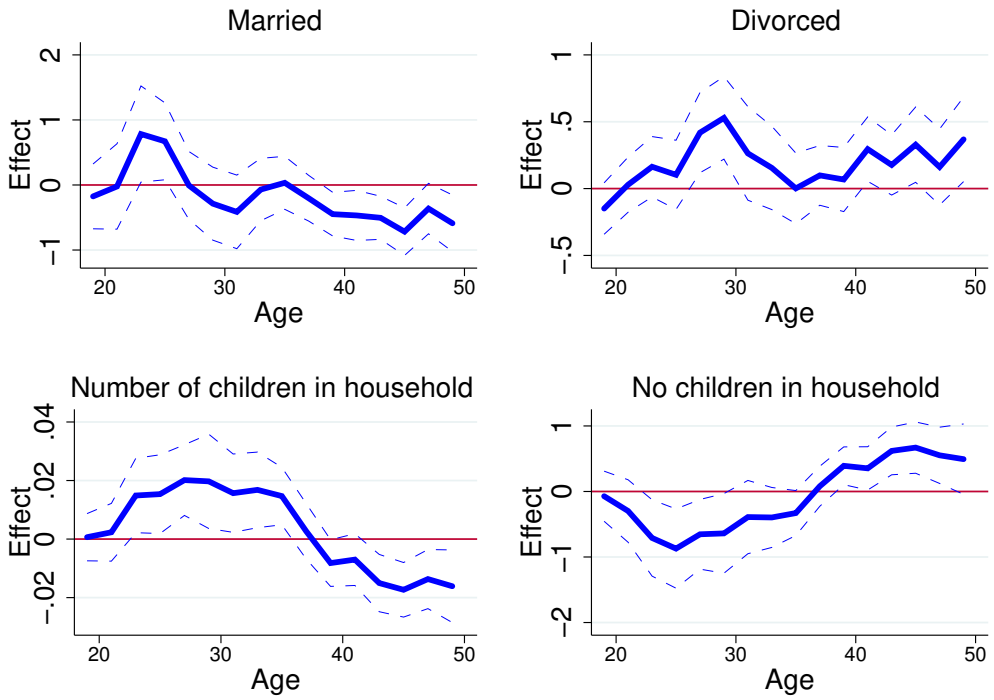


Notes: Results are based on the Mincerian specification (Eq. (2)) and double-weighted specification (Eq. (3)), using data from the ASEC Supplement to CPS from 1976 to 2016, the Censuses from 1980, 1990, 2000, and the ACS from 2001–2016. "SvW" refers to the specification estimated in Schwandt and von Wachter (2019).

Figure 7: Effect of State Unemployment Rate at Labor Market Entry on Economic Outcomes

Coefficients and 95 percent confidence intervals are plotted from regressions of economic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Figure 8: Effect of State Unemployment Rate at Labor Market Entry on Family Formation Outcomes



Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of family formation outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Table 1: Summary Statistics

Panel A: Mortality (deaths per 10,000)

		Disease-related deaths				External causes				
	Total	Total	Heart	Lung cancer	Liver	Total	Acci-dents	Suicide	Drug pois.	Deaths of despair
Age										
19-25	11.32	2.73	0.51	0.02	0.03	8.59	5.16	1.53	0.20	1.76
26-30	12.71	5.79	2.38	0.04	0.12	6.91	3.16	1.56	0.47	2.15
31-35	14.94	8.79	3.37	0.13	0.26	6.15	2.55	1.53	0.86	2.65
36-40	18.57	11.99	2.79	0.37	0.55	6.58	2.50	1.62	1.55	3.73
41-45	26.62	19.06	4.60	0.97	1.07	7.55	2.55	1.86	2.41	5.35
46-49	36.99	28.70	7.13	1.87	1.79	8.29	2.51	2.14	2.99	6.93

Panel B: Socioeconomic Outcomes

	Economic outcomes				Family formation outcomes			
	Log earnings	Any earnings (percent)	Any welfare (percent)	In poverty (percent)	Married (percent)	Divorced (percent)	Number children	No children (percent)
Age								
19-29	9.31	82.52	5.30	13.93	37.83	6.29	0.51	69.67
30-39	10.03	81.61	3.49	9.59	63.49	14.90	1.35	35.40
40-49	10.20	78.20	1.17	8.99	63.56	19.34	1.20	38.72

Notes: Death rates in panel A are based on Vital Statistics, Census, and ACS data. Panel B shows socioeconomic outcomes from the Census, ACS, and CPS, aggregated using population weights. Sample in both panels includes cohorts born in the US from 1960 through 1967 who are observed from age 19 to age 49. "Deaths of despair" are the sum of "Liver," "Suicide," and "Drug poisoning." "Heart" and "Liver" refer to deaths due to heart and liver disease, respectively. "Drug poisoning" refers to accidental drug poisonings. "Number children" refers to the number of own children living in the household. "No children" refers to respondents without any own children living in the household.

Table 2: Effect of State Unemployment Rate at Labor Market Entry on Mortality for Different Birth Cohorts

	Dep. var.: Deaths per 10,000				
	(1)	(2)	(3)	(4)	(5)
	Cohort '60-'91	Cohort '60-'86	Cohort '60-'76	Cohort '60-'71	Cohort '60-'67
Age 19	-0.100**	-0.101**	-0.097	-0.201**	-0.242**
	(0.049)	(0.050)	(0.068)	(0.079)	(0.104)
Age 20-25	-0.016	-0.083**	-0.028	-0.040	-0.095
	(0.043)	(0.038)	(0.047)	(0.058)	(0.080)
Age 26-30		0.002	-0.037	-0.107**	-0.038
		(0.042)	(0.041)	(0.046)	(0.057)
Age 31-35			0.070	0.022	-0.004
			(0.053)	(0.051)	(0.057)
Age 36-40			0.195***	0.149***	0.111**
			(0.059)	(0.057)	(0.053)
Age 41-45				0.485***	0.415***
				(0.091)	(0.091)
Age 46-49					0.805***
					(0.146)
Total deaths	811,140	1,208,897	1,703,722	1,839,011	1,730,886
Mean d. v.	10.41	11.00	13.35	15.89	18.96
R^2	0.66	0.67	0.81	0.89	0.92

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: Regressions of death rates on the double-weighted labor market entry unemployment rate interacted with age group dummies (Eq. (3)) are shown. Data from Vital Stats, Census, and ACS, years 1979–2016. Each column shows one regression, varying age ranges and restricting cohorts to those observed across the entire analyzed age range. Total deaths vary accordingly. Regressions include fixed effects for age groups, cohort, year, and state of birth. Standard errors (in parenthesis) clustered at the state-of-birth by year-of-birth level. Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Observations population-weighted.

Table 3: Effect of State Unemployment Rate at Labor Market Entry on Disease-Related Causes of Death (Log Death Rate)

	(1)	(2)	(3)	(4)	(5)
	Disease-related	Liver disease	Heart disease	Lung cancer	Other cancer
Age 19	0.006	-0.168	-0.032	-0.099	-0.005
	(0.013)	(0.114)	(0.029)	(0.090)	(0.022)
Age 20-25	0.010**	-0.007	-0.017*	-0.071**	-0.018*
	(0.005)	(0.034)	(0.009)	(0.036)	(0.009)
Age 26-30	0.009*	0.020	-0.022**	-0.066**	-0.006
	(0.005)	(0.023)	(0.009)	(0.029)	(0.008)
Age 31-35	0.018***	0.015	0.022***	-0.026	-0.013*
	(0.004)	(0.019)	(0.008)	(0.017)	(0.007)
Age 36-40	0.020***	0.044***	0.043***	0.002	-0.001
	(0.004)	(0.014)	(0.007)	(0.015)	(0.006)
Age 41-45	0.029***	0.038***	0.043***	0.052***	0.007
	(0.005)	(0.012)	(0.007)	(0.015)	(0.006)
Age 46-49	0.031***	0.057***	0.051***	0.072***	0.007
	(0.005)	(0.012)	(0.007)	(0.014)	(0.005)
Total deaths	1,055,846	50,354	285,436	43,891	222,228
Mean dep. var.	2.11	-.828	.744	-.967	.538
R^2	0.95	0.83	0.89	0.87	0.87

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: Regressions of cause-specific death rates on the double-weighted labor market entry unemployment rate interacted with age group dummies (Eq. (3)) are shown. Death rates in logs as disease-related deaths increase sharply with age. Data from Vital Stats, Census, and ACS, years 1979–2016. Regressions include fixed effects for age groups, cohort, year, and state of birth. Standard errors (in parenthesis) clustered at the state-of-birth by year-of-birth level. Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Observations population-weighted.

Table 4: Effect on Mortality by Demographic Group

Panel A: Overall and disease-related deaths

	Log death rate				Log disease-related mortality			
	White males	White females	Non-white males	Non-white females	White males	White females	Non-white males	Non-white females
Age 19-29	-0.003	0.012**	0.022***	0.024***	-0.001	0.006	0.028**	0.009
	(0.004)	(0.006)	(0.008)	(0.008)	(0.007)	(0.007)	(0.013)	(0.013)
Age 30-39	0.005	0.015***	0.014*	0.026***	0.021***	0.011*	0.038***	0.030***
	(0.003)	(0.005)	(0.007)	(0.008)	(0.005)	(0.006)	(0.010)	(0.010)
Age 40-49	0.016***	0.031***	0.011*	0.036***	0.031***	0.034***	0.046***	0.048***
	(0.004)	(0.006)	(0.007)	(0.008)	(0.006)	(0.007)	(0.010)	(0.011)
Total deaths	806,239	411,691	327,306	185,650	429,244	280,665	198,773	147,164
Mean d.v.	3.01	2.22	3.46	2.7	2.14	1.67	2.75	2.36
R^2	0.86	0.88	0.72	0.81	0.91	0.89	0.86	0.85

Panel B: Deaths of despair

	Log deaths of despair				Log despair + lung cancer + heart disease			
	White males	White females	Non-white males	Non-white females	White males	White females	Non-white males	Non-white females
Age 19-29	-0.015*	-0.026*	0.035**	0.050**	-0.013**	-0.022*	0.018	0.010
	(0.008)	(0.014)	(0.014)	(0.024)	(0.006)	(0.012)	(0.015)	(0.018)
Age 30-39	-0.011	0.004	0.022	0.035	0.003	0.021*	0.031***	0.028*
	(0.007)	(0.014)	(0.013)	(0.021)	(0.005)	(0.011)	(0.012)	(0.015)
Age 40-49	0.015**	0.051***	0.004	0.020	0.027***	0.062***	0.055***	0.085***
	(0.007)	(0.013)	(0.014)	(0.017)	(0.005)	(0.011)	(0.010)	(0.015)
Total deaths	187,558	74,239	41,592	16,505	343,319	138,365	113,868	53,669
Mean d.v.	1.51	.416	1.34	.438	2.04	.953	2.27	1.41
R^2	0.69	0.76	0.54	0.58	0.85	0.84	0.77	0.73

Notes: Data and specification as in Table 3. White refers to non-Hispanic white. Non-white refers to non-white and Hispanic white. p-value of t-test for race difference of age 40-49 coefficient for Log deaths of despair is 0.075. All other age 40-49 coefficients not significantly different between whites and non-whites. Deaths of despair include liver disease, drug poisoning, and suicide. Mean cause-specific mortality rates by age and demographic group are reported in Table A.1.

Table 5: Predicted Impacts of Effect ofUnemployment Rate at Labor Market Entry On Overall Cohort Mortality

	1982 recession graduation cohort (+3.9 p.p. unemployment rate)	
	linear extrapolation (1)	percent extrapolation (2)
Excess deaths (in cohort of 4 million)		
At age 0	0	0
At age 20	-374	-374
At age 50	7,089	7,089
At age 65	41,741	39,183
At age 80	90,624	121,327
Cohort share deceased prematurely		
At age 65	1.04	0.98
At age 80	2.27	3.03
Loss in life expectancy at birth		
Months per capita	-5.90	-8.91
Sum of life years (in cohort of 4m)	-1,973,297	-2,978,912

Notes: Linear extrapolation assumes that after age 50 the mortality rate increases annually by an additional 0.09 per 10,000 and percent extrapolation assumes that after age 50 the mortality rate is increased by a constant 2 percent, respectively, for each percentage point increase in the graduation year unemployment rate. Excess deaths and life years lost are based on initial birth cohort sizes of 4 million (the 1964 birth cohort is 4,027,490). Mortality effects prior to age 50 are taken from the baseline results shown in Figure 2. Appendix Figure A.6 plots the extrapolated effect patterns.

Table 6: Effect of State Unemployment Rate at Labor Market Entry on Socioeconomic Outcomes

Panel A: Economic Outcomes

	Log earnings	Any earnings (percent)	Any welfare (percent)	In poverty (percent)
Age 19-29	-0.019***	-0.260*	0.491***	0.445***
	(0.004)	(0.133)	(0.055)	(0.093)
Age 30-39	-0.007**	0.363***	0.055	0.026
	(0.003)	(0.130)	(0.049)	(0.088)
Age 40-49	-0.009***	0.239**	-0.154***	0.056
	(0.003)	(0.118)	(0.040)	(0.073)
Observations	6,805,719	6,807,383	6,807,383	6,807,377
Mean dep. var.	9.9	80.4	3.01	10.6
R^2	0.77	0.27	0.25	0.33

Panel B: Family Formation Outcomes

	Married (percent)	Divorced (percent)	Number of children	No children (percent)
Age 19-29	0.219	0.165**	0.012***	-0.556***
	(0.234)	(0.083)	(0.005)	(0.203)
Age 30-39	-0.208	0.132	0.009**	-0.159
	(0.138)	(0.086)	(0.004)	(0.126)
Age 40-49	-0.499***	0.229***	-0.013***	0.527***
	(0.147)	(0.082)	(0.004)	(0.160)
Observations	6,807,383	6,807,383	6,807,383	6,807,383
Mean dep. var.	56.3	14.4	1.05	46.5
R^2	0.61	0.37	0.65	0.68

Notes: Regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with age group dummies (Eq. (3)) are shown. Data from Census, ACS, and CPS, years 1979–2016. Regressions include fixed effects for age groups, cohort, year, and state of birth. Standard errors (in parenthesis) clustered at the state-of-birth by year-of-birth level. Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Observations weighted by population-weighted cell size.

Table 7: Effect on Economic Outcomes by Demographic Group

Panel A: Labor market outcomes

	Log earnings				Any earnings (percent)			
	White males	White females	Non-white males	Non-white females	White males	White females	Non-white males	Non-white females
Age 19-29	-0.019***	-0.008	-0.040***	-0.041***	-0.015	-0.461**	-0.027	-1.294***
	(0.005)	(0.007)	(0.009)	(0.010)	(0.174)	(0.180)	(0.348)	(0.374)
Age 30-39	-0.010**	-0.006	-0.005	-0.002	0.190	0.420**	0.674**	0.259
	(0.004)	(0.006)	(0.008)	(0.008)	(0.165)	(0.193)	(0.277)	(0.306)
Age 40-49	-0.021***	-0.005	0.006	0.004	-0.117	0.397**	0.717***	0.371
	(0.004)	(0.005)	(0.007)	(0.007)	(0.168)	(0.166)	(0.259)	(0.289)
Obs.	2,706,863	2,781,916	617,265	699,675	2,706,875	2,781,979	617,845	700,684
Mean d.v.	10.2	9.7	9.84	9.54	86.6	77.1	79.9	71.8
t-test		0.008	0.001	0.002		0.011	0.007	0.147

Panel B: Welfare and poverty

	Any welfare income (percent)				Below poverty threshold (percent)			
	White males	White females	Non-white males	Non-white females	White males	White females	Non-white males	Non-white females
Age 19-29	0.244***	0.366***	0.698***	1.862***	0.145	0.442***	0.567*	1.726***
	(0.057)	(0.086)	(0.161)	(0.294)	(0.115)	(0.150)	(0.326)	(0.326)
Age 30-39	0.057	0.044	0.031	0.245	0.028	0.142	-0.151	-0.138
	(0.047)	(0.074)	(0.142)	(0.248)	(0.107)	(0.129)	(0.294)	(0.273)
Age 40-49	-0.058	-0.160***	-0.181*	-0.587***	0.065	0.152	-0.122	-0.163
	(0.039)	(0.061)	(0.099)	(0.189)	(0.104)	(0.113)	(0.279)	(0.234)
Obs.	2,706,875	2,781,979	617,845	700,684	2,706,875	2,781,979	617,840	700,683
Mean d.v.	1.11	2.85	2.45	10	6.59	9.05	15.4	23.5
t-test		0.134	0.256	0.006		0.562	0.512	0.389

Notes: Data and specification as in Table 6. White refers to non-Hispanic white. Non-white refers to non-white and Hispanic white. The t-test row reports the p-value of the difference between the age 40-49 coefficient in the each column with the the same coefficient in the regressions for whites. Mean socioeconomic outcomes by age and demographic group are reported in Table A.2.

Table 8: Effect of State Unemployment Rate at Labor Market Entry on Family Formation Outcomes by Demographic Group

Panel A: Marital status

	Married (percent)				Divorced (percent)			
	White males	White females	Non-white males	Non-white females	White males	White females	Non-white males	Non-white females
Age 19-29	0.456*	0.489*	-0.208	-0.428	0.326**	0.218	0.096	-0.338
	(0.261)	(0.286)	(0.438)	(0.419)	(0.130)	(0.150)	(0.184)	(0.235)
Age 30-39	-0.309	-0.277	-0.320	0.301	0.490***	0.016	0.291	-0.602**
	(0.196)	(0.200)	(0.424)	(0.398)	(0.147)	(0.147)	(0.297)	(0.282)
Age 40-49	-0.770***	-0.629***	-0.211	0.308	0.564***	0.079	0.329	-0.121
	(0.200)	(0.193)	(0.353)	(0.306)	(0.152)	(0.138)	(0.205)	(0.210)
Obs.	2,706,875	2,781,979	617,845	700,684	2,706,875	2,781,979	617,845	700,684
Mean d.v.	57.3	63.2	43.1	42.6	12.2	15.3	13.2	19
t-test		0.588	0.179	0.004		0.022	0.360	0.014

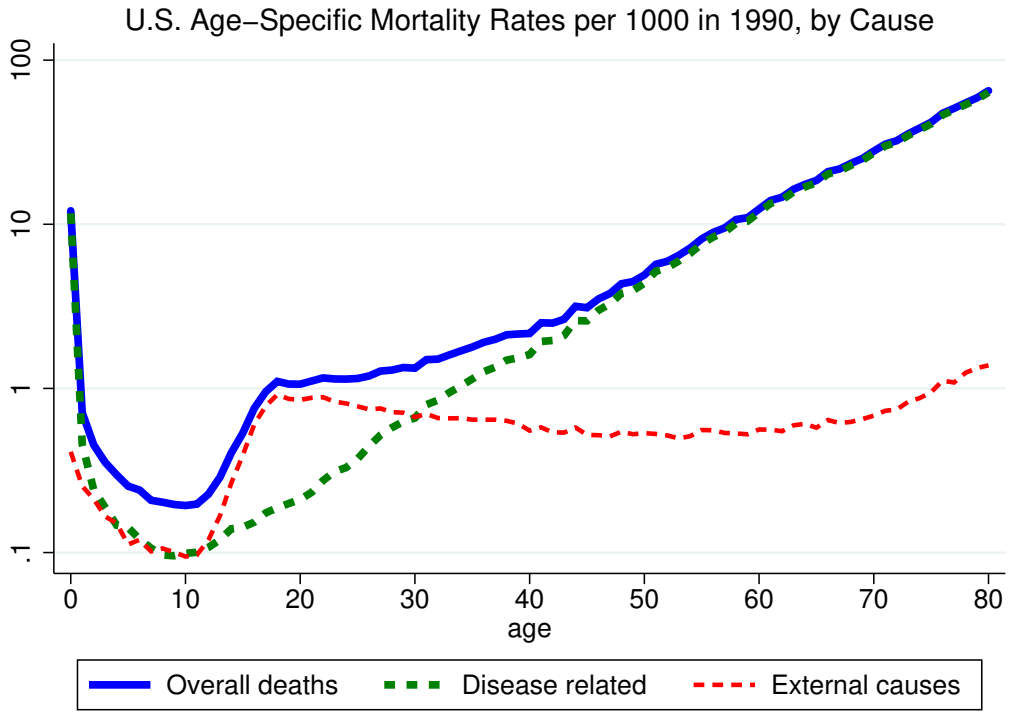
Panel B: Own children in household

	Number of children				No children (percent)			
	White males	White females	Non-white males	Non-white females	White males	White females	Non-white males	Non-white females
Age 19-29	0.008	0.024***	-0.004	0.011	-0.278	-1.129***	-0.146	-0.562
	(0.005)	(0.006)	(0.009)	(0.009)	(0.215)	(0.272)	(0.358)	(0.370)
Age 30-39	0.006	0.013**	-0.012	0.006	0.096	-0.347*	0.206	-0.304
	(0.005)	(0.005)	(0.011)	(0.011)	(0.194)	(0.204)	(0.403)	(0.368)
Age 40-49	-0.014***	-0.018***	-0.009	-0.015	0.762***	0.656***	0.145	0.121
	(0.005)	(0.005)	(0.009)	(0.009)	(0.190)	(0.218)	(0.365)	(0.341)
Obs.	2,706,875	2,781,979	617,845	700,684	2,706,875	2,781,979	617,845	700,684
Mean d.v.	.948	1.16	.788	1.25	52.1	39.9	60.6	37.1
t-test		0.584	0.599	0.924		0.651	0.123	0.091

Notes: Data and specification as in Table 6. White refers to non-Hispanic white. Non-white refers to non-white and Hispanic white. The t-test row reports the p-value of the difference between the age 40-49 coefficient in the each column with the the same coefficient in the regressions for whites. Mean socioeconomic outcomes by age and demographic group are reported in Table A.2.

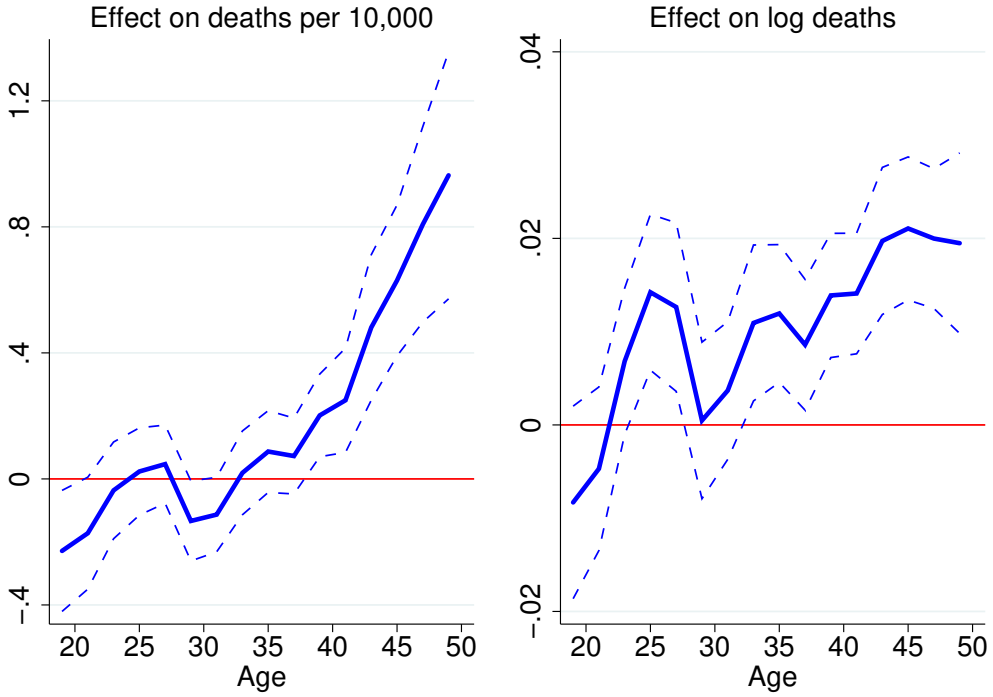
Appendix

Figure A.1: Average Mortality by Age (log scale)



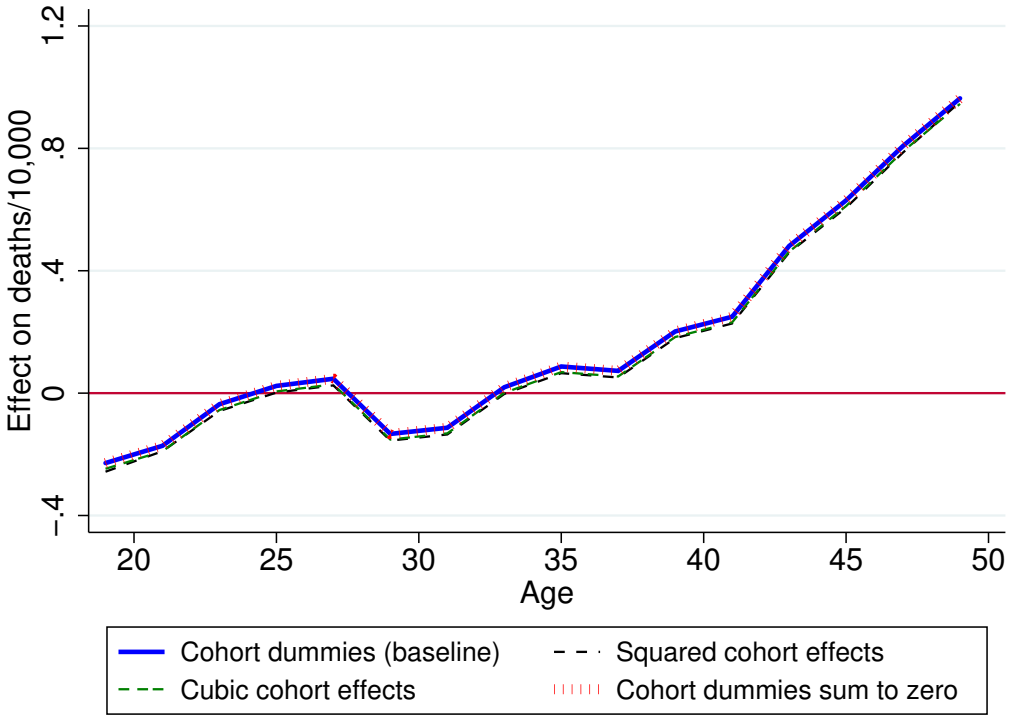
Notes: Lifecycle mortality profiles are shown by cause, based on Vital Stats and Census data for 2000. External causes include all deaths that are not disease-related.

Figure A.2: Effect of Labor Market Entry Unemployment Rate on Mortality



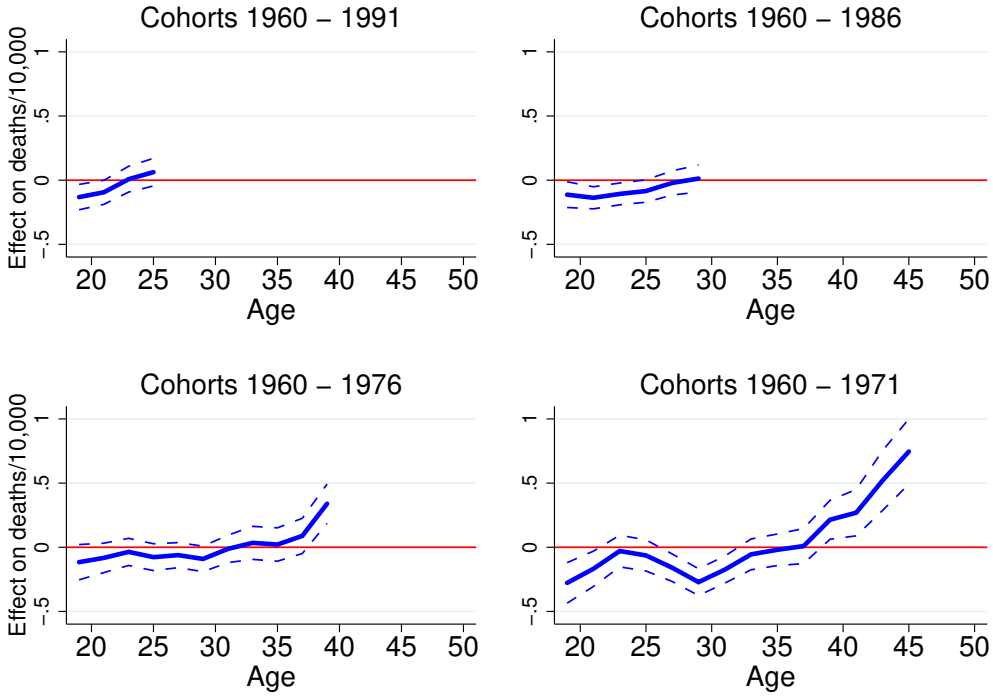
Notes: This figure shows baseline results regressing death rates in levels (left, replicating Figure 2) and in logs (right).

Figure A.3: Baseline Mortality Results With Alternative Cohort Effects



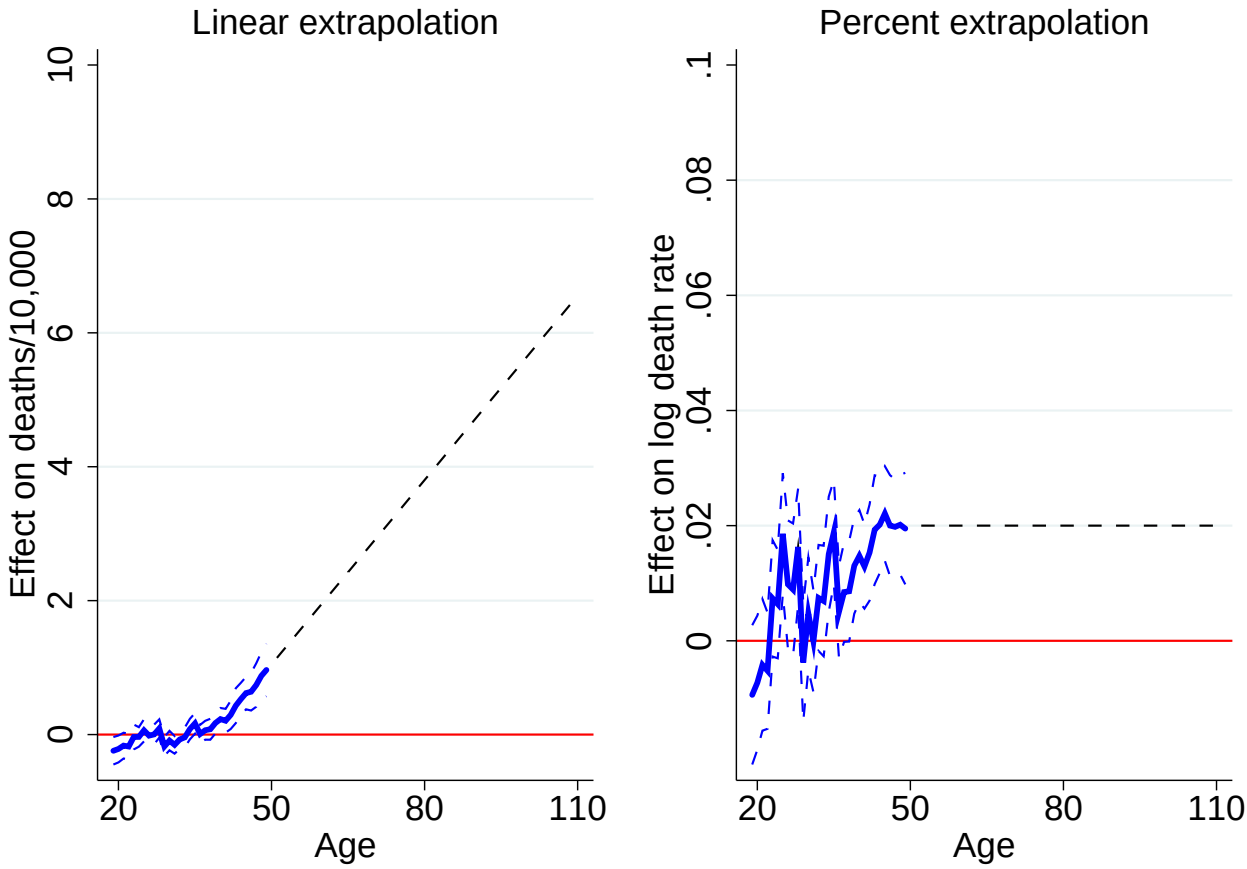
Notes: The blue solid line replicates the baseline results as shown in Figure 2. The dashed lines show results from specifications that replace cohort fixed effects with squared and cubic cohort effects. The red dashed line shows coefficients from a specification in which we restrict cohort effects to sum to zero (as suggested by Deaton (1997)).

Figure A.4: Effect of Labor Market Entry Unemployment Rate on Mortality, by Cohort



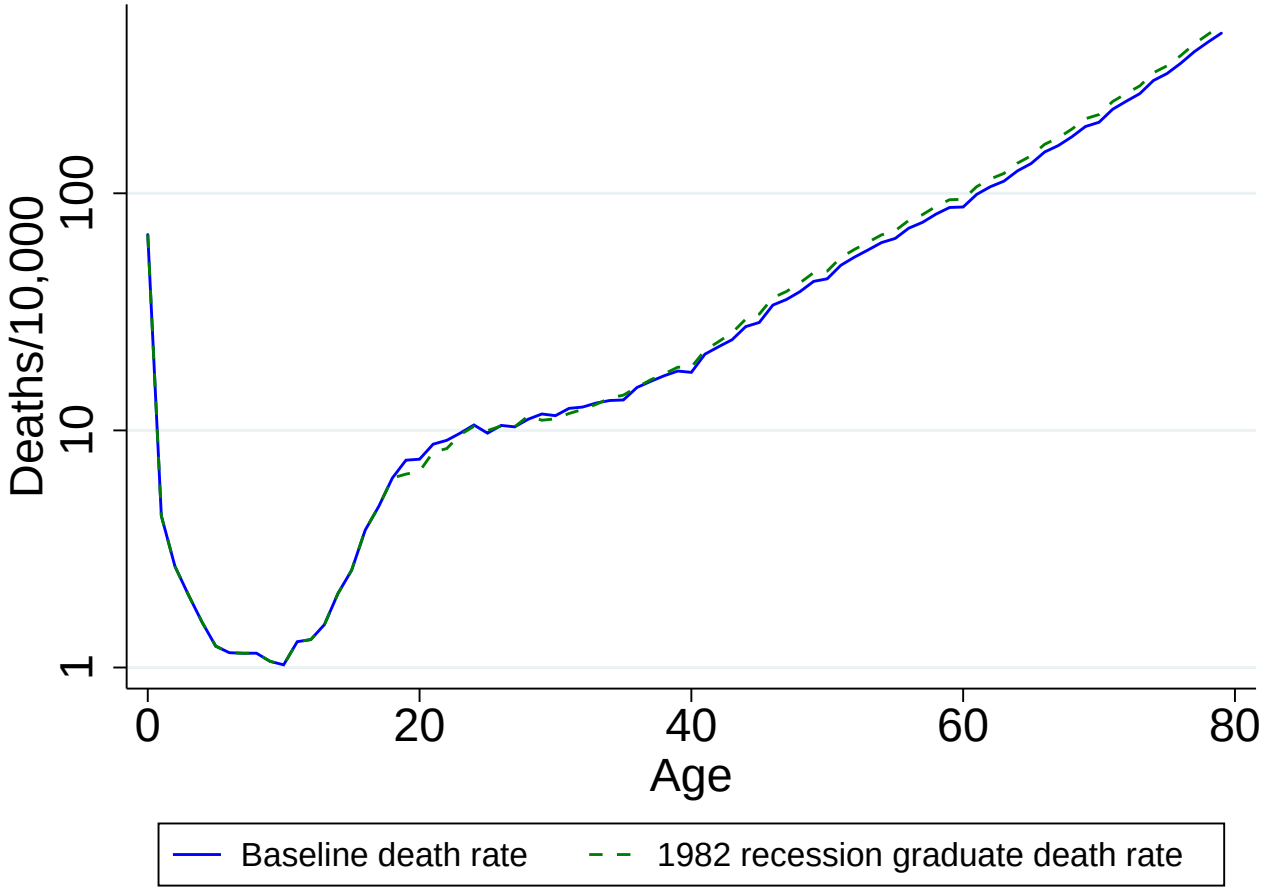
Notes: This figure plots coefficients of the baseline specification for different age/cohort ranges for 2-year age groups. Numerical results for 5-year age groups are reported in Table 2.

Figure A.5: Extrapolating Effects into Old Age



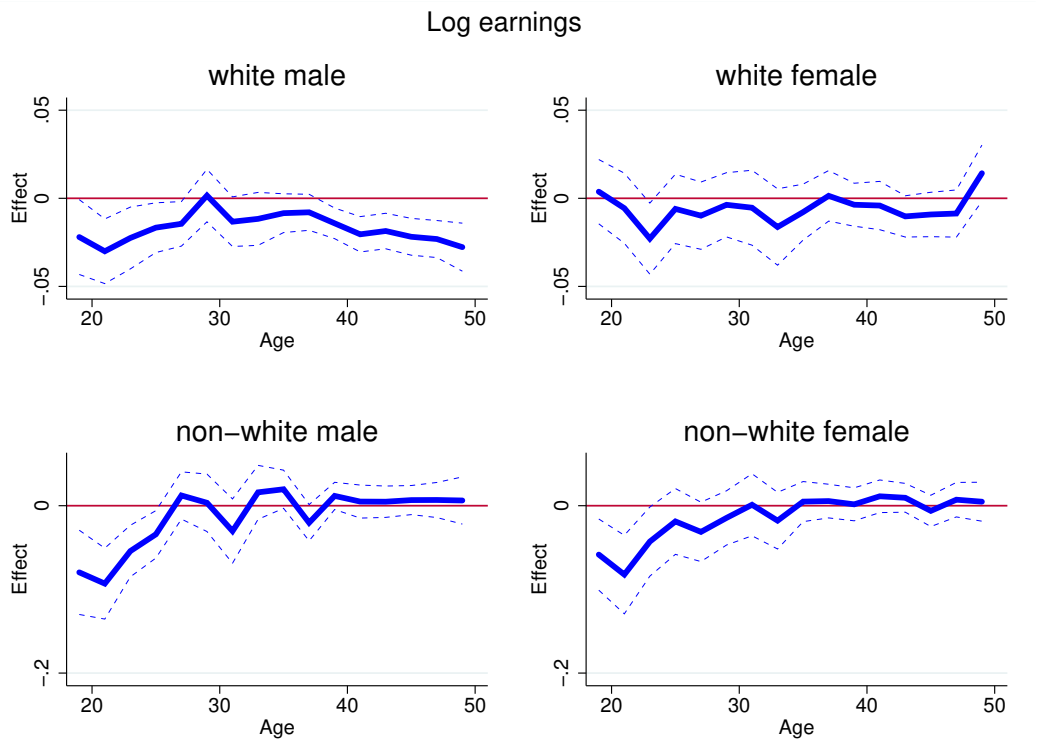
Notes: This figure shows linear and percent extrapolations used in Table 5.

Figure A.6: Mortality Profile of Baseline and 1982 Graduate Death Rate (+3.9 Percentage Point Entry Unemployment Rate), Using a Constant 2 Percent Extrapolation



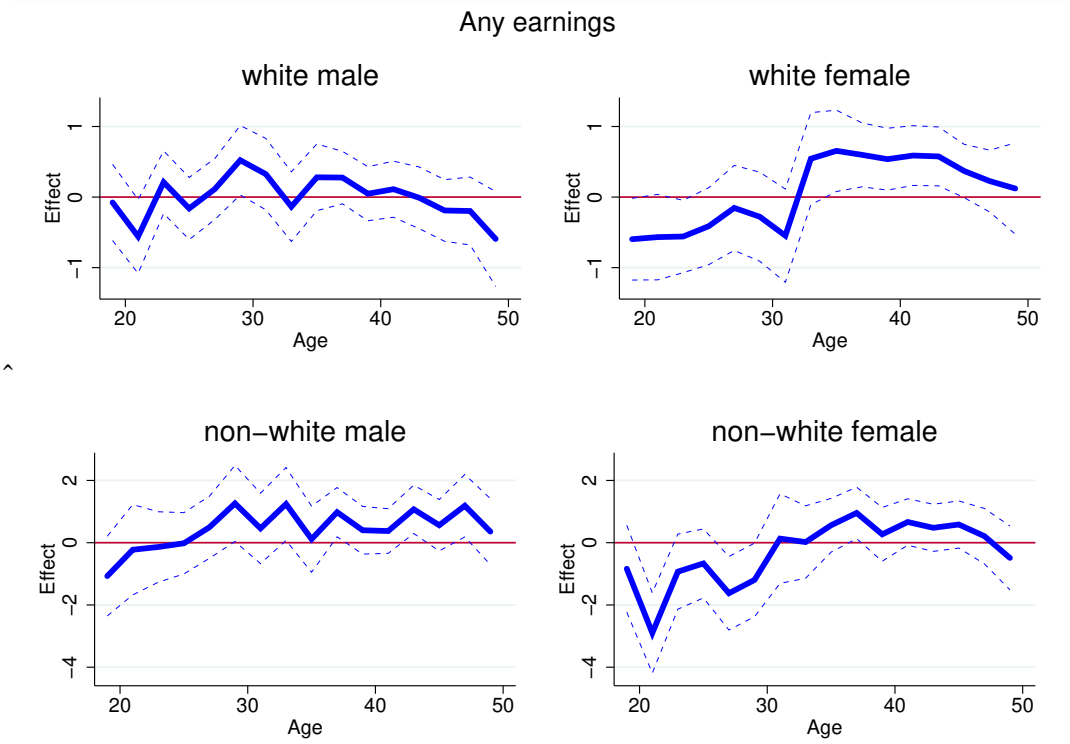
Notes: Baseline death rate calculated based on cross-sectional mortality data for 2010. The Great Recession–graduate death rate is constructed by adding the effect of a five percentage point increase in the graduation year unemployment rate. A constant 2 percent effect is assumed for ages above age 50 (i.e. 10 percent is added to the age specific mortality rate).

Figure A.7: Effect of Labor Market Entry Unemployment Rate on Earnings by Demographic Group



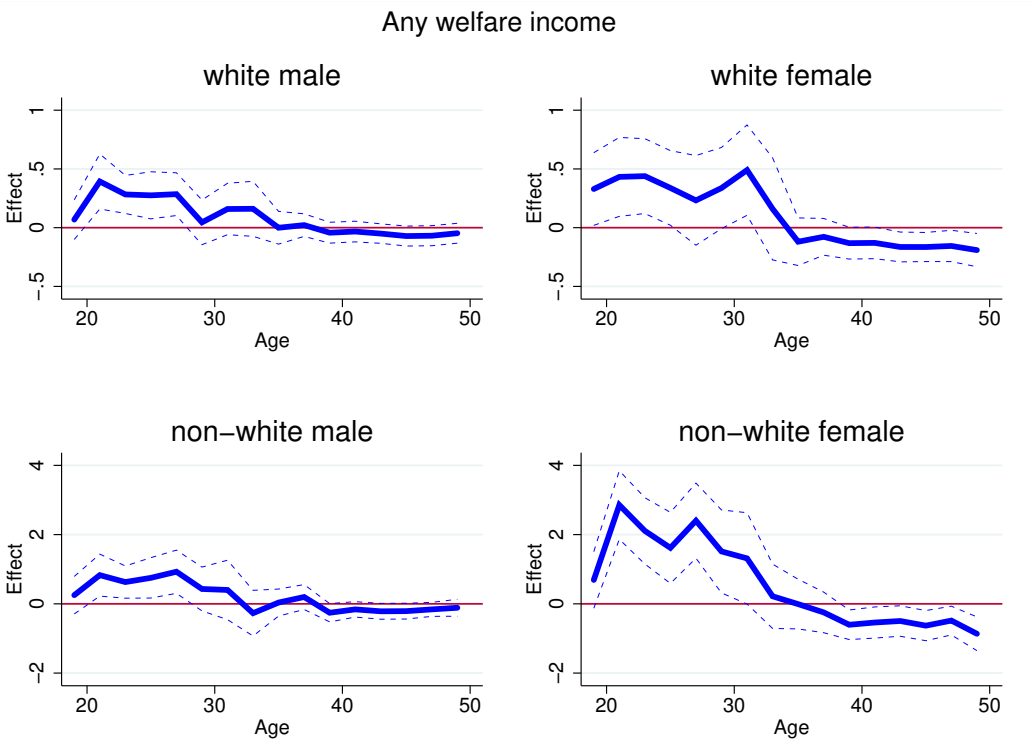
Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Figure A.8: Effect of Labor Market Entry Unemployment Rate on Percent with Any Earnings by Demographic Group



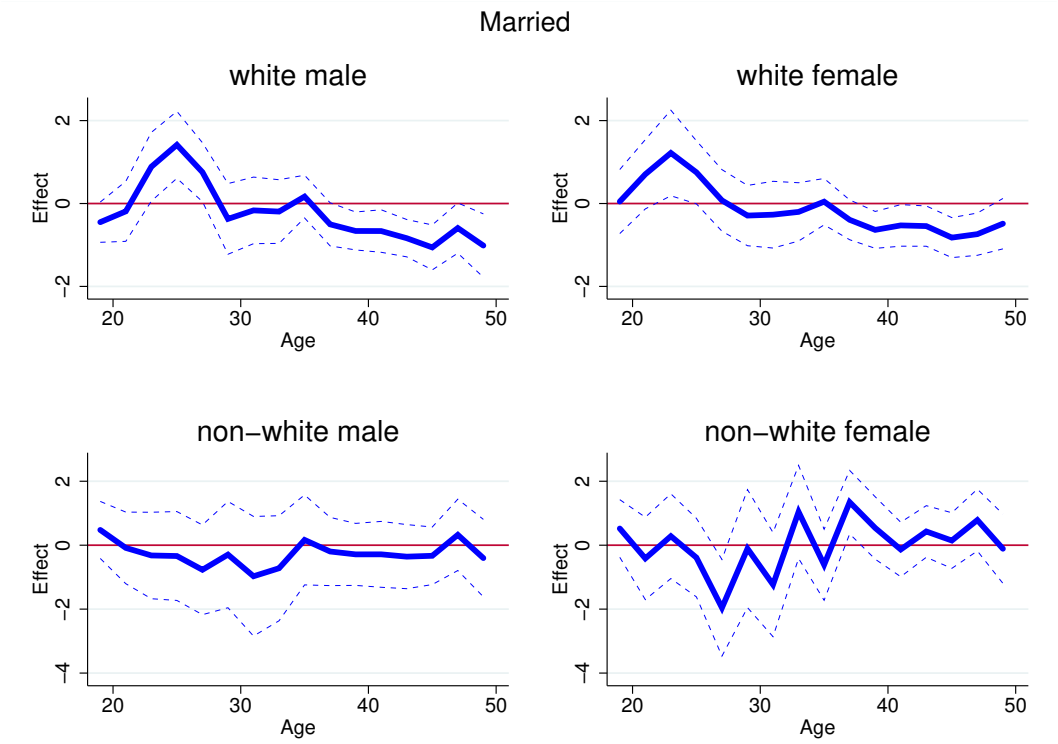
Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Figure A.9: Effect of Labor Market Entry Unemployment Rate on Welfare Receipt by Demographic Group



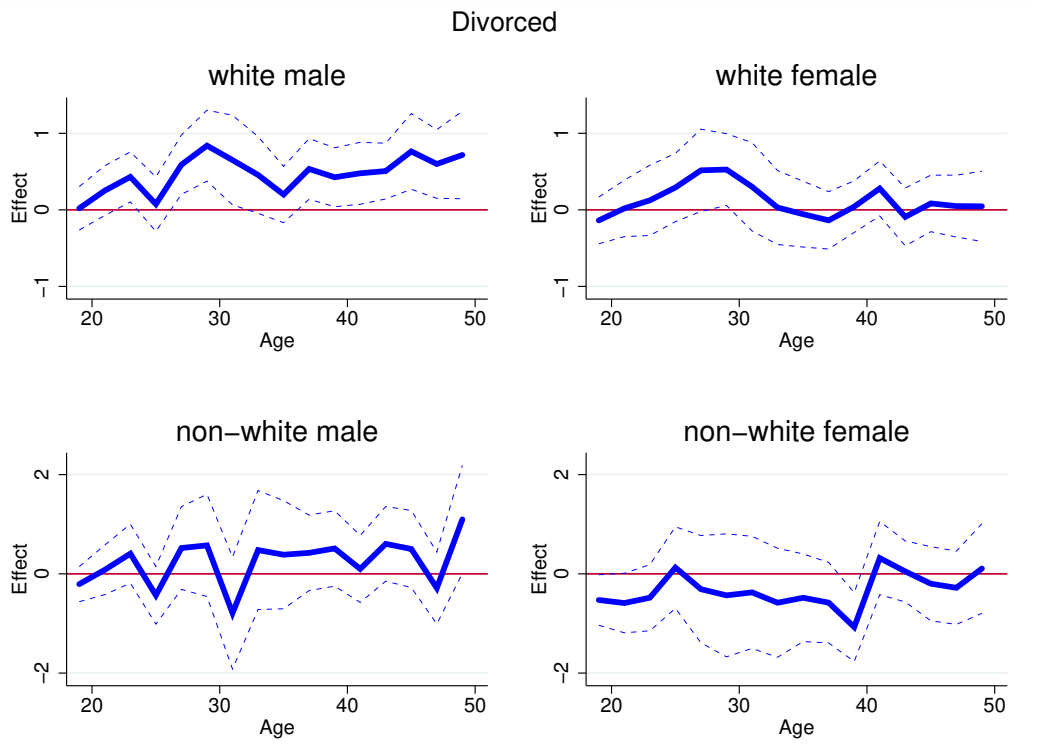
Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Figure A.10: Effect of Labor Market Entry Unemployment Rate on Percent Married by Demographic Group



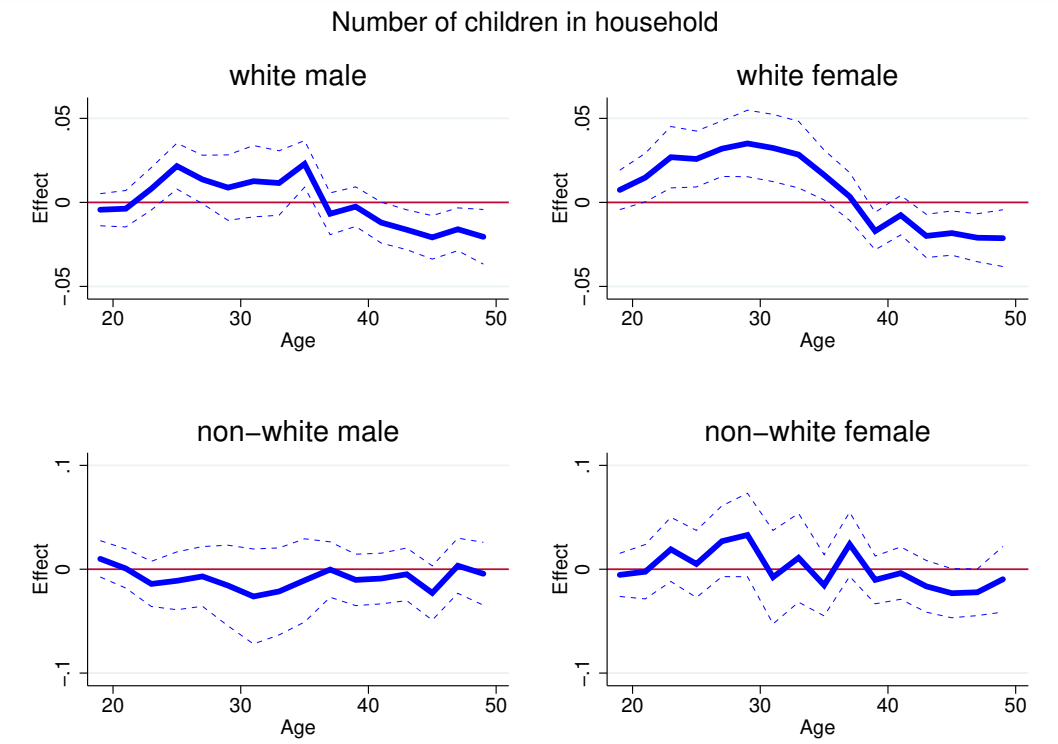
Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Figure A.11: Effect of Labor Market Entry Unemployment Rate on Percent Divorced by Demographic Group



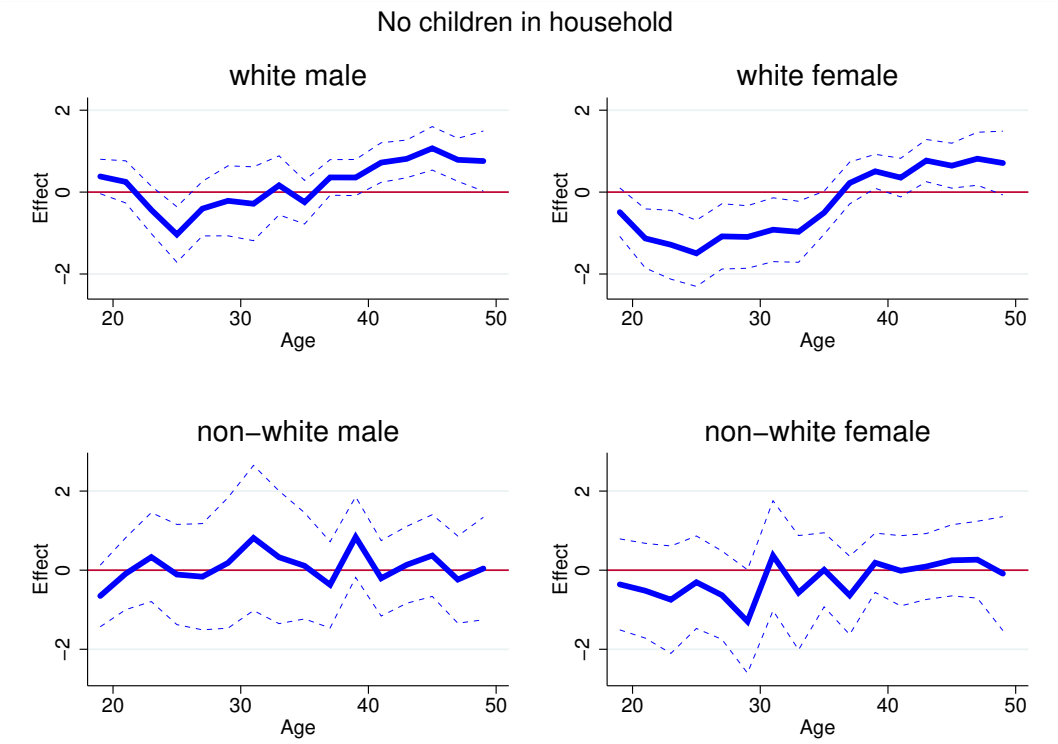
Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Figure A.12: Effect of Labor Market Entry Unemployment Rate on the Number of Own Children in Household by Demographic Group



Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Figure A.13: Effect of Labor Market Entry Unemployment Rate on Percent Without Own Children in household by Demographic Group



Notes: Coefficients and 95 percent confidence intervals are plotted from regressions of socioeconomic outcomes on the double-weighted labor market entry unemployment rate interacted with 2-year age group dummies (Eq. (3)), including fixed effects for age groups, cohort, year, and state of birth. Data from Census, ACS, and CPS, years 1979–2016. Further regression and sample details as in Table 2.

Table A.1: Cause-Specific Deaths per 10,000, by Age Group and Demographic Group

		Disease-related deaths				External causes				
	Total	Total	Heart	Lung cancer	Liver	Total	Acci-dents	Suicide	Drug pois.	Deaths of despair
White male										
19-25	16.09	2.93	0.54	0.02	0.02	13.16	8.83	2.79	0.26	3.07
26-30	15.45	6.17	2.85	0.04	0.11	9.28	4.85	2.74	0.61	3.46
31-35	17.65	9.06	3.96	0.15	0.29	8.59	3.81	2.69	1.18	4.16
36-40	21.37	11.96	3.42	0.37	0.65	9.41	3.68	2.86	2.09	5.60
41-45	30.92	20.00	6.05	1.00	1.35	10.92	3.86	3.27	3.13	7.75
46-49	43.35	31.41	9.52	1.97	2.31	11.93	3.81	3.78	3.70	9.79
White female										
19-25	5.44	1.87	0.24	0.01	0.01	3.57	2.38	0.56	0.11	0.69
26-30	5.62	2.94	0.55	0.03	0.06	2.68	1.47	0.57	0.20	0.83
31-35	7.61	4.87	0.92	0.13	0.16	2.75	1.27	0.67	0.41	1.24
36-40	11.78	8.19	1.30	0.37	0.38	3.59	1.38	0.82	1.03	2.24
41-45	18.47	13.73	2.29	0.97	0.72	4.74	1.38	1.08	1.97	3.78
46-49	27.22	21.60	3.71	1.87	1.27	5.62	1.36	1.30	2.67	5.24
Non-white male										
19-25	20.77	4.63	1.26	0.02	0.05	16.15	5.90	1.92	0.35	2.32
26-30	31.34	13.96	7.48	0.06	0.27	17.38	5.12	2.36	1.11	3.74
31-35	33.37	20.09	10.36	0.16	0.50	13.28	4.10	1.96	1.76	4.23
36-40	34.09	22.59	6.09	0.40	0.93	11.50	3.87	1.64	2.43	5.01
41-45	43.87	32.63	8.72	1.09	1.63	11.25	3.92	1.52	3.00	6.15
46-49	54.65	44.04	12.83	1.91	2.45	10.61	3.56	1.45	3.38	7.29
Non-white female										
19-25	6.95	3.28	0.61	0.01	0.06	3.66	1.46	0.31	0.16	0.53
26-30	11.73	7.44	2.51	0.04	0.22	4.30	1.43	0.40	0.40	1.02
31-35	15.14	11.37	3.81	0.10	0.32	3.77	1.42	0.34	0.63	1.29
36-40	19.88	16.18	3.08	0.35	0.53	3.70	1.32	0.38	1.04	1.95
41-45	27.26	23.33	4.39	0.86	1.00	3.92	1.26	0.40	1.56	2.96
46-49	36.37	32.13	6.46	1.62	1.48	4.25	1.30	0.40	1.98	3.87

Notes: Death rates based on Vital Statistics, Census, and ACS data. Sample includes cohorts born in the US from 1960 through 1967 who are observed from age 19 to age 49. "Deaths of despair" are the sum of "Liver," "Suicide," and "Drug poisoning." "Heart" and "Liver" refer to deaths due to heart and liver disease, respectively. "Drug poisoning" refers to accidental drug poisonings. White refers to non-Hispanic white. Non-white refers to non-white and Hispanic white.

Table A.2: Economic and Family Formation Outcomes by Age Group and Demographic Group

	Economic outcomes				Family formation outcomes			
	Log earnings	Any earnings (percent)	Any welfare (percent)	In poverty (percent)	Married (percent)	Divorced (percent)	Number children	No children (percent)
White male								
19-29	9.57	90.25	1.62	7.91	33.21	4.47	0.31	80.06
30-39	10.40	88.34	1.36	5.64	65.30	12.51	1.18	42.47
40-49	10.55	83.14	0.62	6.41	66.89	16.87	1.18	41.29
White female								
19-29	9.19	81.40	5.19	11.89	46.39	8.04	0.59	63.76
30-39	9.76	76.08	3.32	8.47	70.56	15.70	1.52	27.00
40-49	9.97	75.20	1.07	7.68	68.83	19.67	1.27	33.73
Non-white male								
19-29	9.28	81.97	3.21	19.30	28.23	3.94	0.37	78.45
30-39	10.04	83.77	3.07	13.06	49.78	14.14	1.03	51.71
40-49	10.16	75.43	1.38	13.90	50.84	20.20	0.97	51.77
Non-white female								
19-29	8.95	65.02	17.63	30.92	35.83	8.64	0.95	49.38
30-39	9.68	75.94	11.36	23.19	46.31	20.78	1.65	25.72
40-49	9.89	74.17	3.21	18.01	45.24	25.89	1.20	35.66

Notes: Data from the Census, ACS, and CPS, aggregated using population weights. Sample in both panels includes cohorts born in the US from 1960 through 1967 who are observed from age 19 to age 49. "Deaths of despair" are the sum of "Liver," "Suicide," and "Drug poisoning." "Heart" and "Liver" refer to deaths due to heart and liver disease, respectively. "Drug poisoning" refers to accidental drug poisonings. "Number children" refers to the number of own children living in the household. "No children" refers to respondents without any own children living in the household. White refers to non-Hispanic white. Non-white refers to non-white and Hispanic white.

Table A.3: Effect of Labor Market Entry Unemployment Rates on Log Mortality for Different Birth Cohorts

	Dep. var.: Log death rate				
	(1)	(2)	(3)	(4)	(5)
	Cohort '60-'91	Cohort '60-'86	Cohort '60-'76	Cohort '60-'71	Cohort '60-'67
Age 19	-0.011**	-0.008*	-0.007	-0.008	-0.009
	(0.005)	(0.004)	(0.005)	(0.005)	(0.006)
Age 20-25	-0.002	-0.006*	0.002	0.006**	0.003
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)
Age 26-30		0.003	0.004	0.004	0.007*
		(0.004)	(0.003)	(0.003)	(0.004)
Age 31-35			0.005	0.007**	0.010***
			(0.003)	(0.003)	(0.003)
Age 36-40			0.004	0.009***	0.010***
			(0.003)	(0.003)	(0.003)
Age 41-45				0.012***	0.018***
				(0.003)	(0.003)
Age 46-49					0.020***
					(0.004)
Total deaths	811,140	1,208,897	1,703,722	1,839,011	1,730,886
Mean d. v.	2.31	2.37	2.54	2.69	2.83
R^2	0.64	0.64	0.80	0.88	0.92

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: Regressions of log death rates on the double-weighted labor market entry unemployment rate interacted with age group dummies (Eq. (3)) are shown. Data from Vital Stats, Census, and ACS, years 1979–2016. Each column shows one regression, varying age ranges and restricting cohorts to those observed across the entire analyzed age range. Total deaths vary accordingly. Regressions include fixed effects for age groups, cohort, year, and state of birth. Standard errors (in parenthesis) clustered at the state-of-birth by year-of-birth level. Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Observations population-weighted.

Table A.4: Effect of Labor Market Entry Unemployment Rate on Deaths of Despair Across Cohorts

	Dep. var.: Deaths per 10,000				
	(1)	(2)	(3)	(4)	(5)
	Cohort '60-'91	Cohort '60-'86	Cohort '60-'76	Cohort '60-'71	Cohort '60-'67
Age 19	0.010	-0.026	-0.022	-0.017	0.003
	(0.016)	(0.016)	(0.020)	(0.021)	(0.024)
Age 20-25	0.026**	-0.008	-0.001	-0.001	-0.004
	(0.013)	(0.013)	(0.016)	(0.016)	(0.018)
Age 26-30		-0.019	-0.013	-0.017	-0.016
		(0.015)	(0.016)	(0.016)	(0.019)
Age 31-35			-0.014	-0.010	-0.021
			(0.017)	(0.017)	(0.018)
Age 36-40			0.043**	0.046**	0.016
			(0.019)	(0.019)	(0.019)
Age 41-45				0.091***	0.081***
				(0.025)	(0.025)
Age 46-49					0.103***
					(0.036)
Total deaths	167,470	254,667	342,565	357,584	319,894
Mean d. v.	1.00	1.00	1.00	1.00	1.00
R^2	0.59	0.63	0.69	0.78	0.83

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: Regressions of log deaths of despair on the double-weighted labor market entry unemployment rate interacted with age group dummies (Eq. (3)) are shown. Data from Vital Stats, Census, and ACS, years 1979–2016. Each column shows one regression, varying age ranges and restricting cohorts to those observed across the entire analyzed age range. Total deaths vary accordingly. Regressions include fixed effects for age groups, cohort, year, and state of birth. Standard errors (in parenthesis) clustered at the state-of-birth by year-of-birth level. Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Observations population-weighted.

Table A.5: Effect on Disease-Related Mortality by Demographic Group, Excluding HIV/Aids

	Log disease-related deaths				Log disease-related, excluding HIV/Aids			
	White males	White females	Non-white males	Non-white females	White males	White females	Non-white males	Non-white females
Age 19-29	-0.001	0.006	0.028**	0.009	-0.009	0.004	0.018	0.002
	(0.007)	(0.007)	(0.013)	(0.013)	(0.007)	(0.008)	(0.015)	(0.013)
Age 30-39	0.021***	0.011*	0.038***	0.030***	0.018***	0.010	0.030***	0.030***
	(0.005)	(0.006)	(0.010)	(0.010)	(0.005)	(0.007)	(0.009)	(0.009)
Age 40-49	0.031***	0.034***	0.046***	0.048***	0.024***	0.029***	0.028***	0.031***
	(0.006)	(0.007)	(0.010)	(0.011)	(0.006)	(0.007)	(0.010)	(0.010)
Total deaths	429,244	280,665	198,773	147,164	385,380	274,793	155,358	129,351
Mean d.v.	2.14	1.67	2.75	2.36	1.97	1.63	2.47	2.22
R^2	0.91	0.89	0.86	0.85	0.91	0.89	0.85	0.85

Notes: White refers to non-Hispanic white. Non-white refers to non-white and Hispanic white. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Mean cause-specific mortality rates by age and demographic group are reported in Table A.1.