NBER WORKING PAPER SERIES

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Working Paper 27512
http://www.nber.org/papers/w27512

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

We thank David Bradford, Stephen Coussens, Leland Farmer, Ellen Meara, Peter Muennig, Sam Norris, Andrew Oswald and seminar participants at Columbia University, the Essen Economics of Mental Health Conference, Florida State University, Luxembourg Institute of Socio-Economic Research, Southern Economic Association meetings, Syracuse University, University of Copenhagen, and University of Hamburg Workshop on Risky Behaviors for helpful comments. Bradley Katcher and Kelsey Pukelis provided superb research assistance. Ruhm is grateful for financial support provided from the University of Virginia Bankard Fund. 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. 27512
July 2020
JEL No. I10,I12,I24,J10

ABSTRACT

We examine whether the least educated population groups experienced the worst mortality trends during the 21st century by measuring changes in mortality across education quartiles. We document sharply differing gender patterns. Among women, mortality trends improved fairly monotonically with education. Conversely, male trends for the lowest three education quartiles were often similar. For both sexes, the gap in average mortality between the top 25 percent and the bottom 75 percent is growing. However, there are many groups for whom these average patterns are reversed – with better experiences for the less educated – or where the differences are statistically indistinguishable.

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1 Introduction

There has long been an interest in the patterns and causes of social disparities in health. Research documenting the "gradient" between health and socioeconomic status dates back at least to the famous Whitehall studies (Marmot, Shipley, and Rose 1984; Marmot et al. 1991), with many analyses conducted since that time.[1] Despite considerable research (Lleras-Muney 2005; Cutler and Lleras-Muney 2010; Clark and Royer 2013), there remain unresolved questions regarding the causes of the inverse relationship between educational attainment and health. Studies have also highlighted differences in mortality trends over time by educational group (Meara, Richards, and Cutler 2008; Cutler et al. 2011; Olshansky et al. 2012). Recently, particular attention has been paid to reversals in historic progress in reducing death rates. Life expectancy at birth in the United States fell from 78.9 to 78.6 years from 2014 to 2017 (National Center for Health Statistics 2017; Murphy et al. 2018), the first such three-year decline in a century. This decrease is part of a broader phenomenon emphasized in the influential research of Case and Deaton (2015), highlighting the increased death rates of middle-aged non-Hispanic whites from 1999-2013.

In this paper, we study whether the least educated have experienced the most adverse trends in mortality since the beginning of the 21st century. A significant challenge in answering this question is that growing disparities may result from secular increases in educational attainment, such that there is increasing negative selection into the lowest schooling categories (Dowd and Hamoudi 2014, Bound et al. 2015). Individuals who, in previous cohorts, would have failed to complete high school may now be high school graduates or even college-educated. These compositional changes could theoretically increase death rates for *all* education groups, even with no overall change. In principle,

[1] In addition to a seminal publication by Kitagawa and Hauser (1973), examples include; Meara et al. (2008), Cutler et al. (2011) and Montez et al. (2011). Bosworth (2018) provides an overview of difficulties in conducting this research, and some recent results.

analyzing mortality at fixed percentiles of the distribution of educational attainment would account for such shifts. However, the information needed to construct death rates by percentile is inconsistently reported across data sources, both in a given year and over time, complicating such efforts.[2]

We make progress on this question by constructing quartiles of educational attainment using population and death data combined with information from Census Bureau surveys. Our distributions of educational attainment are based on single years of schooling for subgroups stratified by sex, race/ethnicity, and five-year age ranges. We then use linear regression to quantify how mortality trends vary from 2001-2017 across education quartiles. To inform our empirical specifications, we first develop a conceptual model demonstrating that equal-sized health shocks may have heterogeneous effects across groups based on differences in health capital. This framework emphasizes potential disparities in the patterns of absolute versus relative mortality changes. We therefore present empirical results for trends in both the level and natural log of death rates. For clarity of exposition, we hereafter use the terms “death rates” or “log death rates” to distinguish between these absolute versus relative changes and “mortality” to indicate more general experiences that are consistent across both measures.

This paper builds on a small number of studies attempting to adjust for changes in the education distribution using different approaches. Bound et al. (2015) employ similar methods to construct quartiles of educational attainment and compare life expectancy between the lowest quartile and the top three combined in 1990 and 2010.[3] They find that the bottom quartile of white women experienced increased death rates, while trends for low-educated white men were flat. The top three quartiles generally experienced large

[2] Rostron et al. (2010) supply a useful discussion of many of these complexities.

[3] They classify quartiles separately by race and sex, while our main specification classifies them separately by sex alone. In Section 5, we also consider quartiles that are both race- and sex-specific.

improvements. Goldring, Lange, and Richards-Shubik (2016) focus on changes in the gradient between education and mortality, rather than changes at specific percentiles, using a non-parametric test. They indicate that the gradient has widened for women but fail to find evidence of this pattern for men. Most recently, Novosad, Rafkin, and Asher (2020) develop a partial identification approach to bound changes in mortality rates between 1992 and 2015. They conclude that the worst mortality experiences are concentrated in the bottom decile of the educational distribution for both whites and blacks. We compare experiences of the lowest decile to those of the remainder of the lower quartile (the 11th to 25th percentiles) in Section 5.4.

Our results provide a more comprehensive picture of recent mortality trends among the full U.S. population than most other studies by including Hispanics and other races in addition to whites and non-Hispanic blacks (hereafter "blacks"), and using data through 2017. Next, we supply a comprehensive analysis of the frequency with which mortality trends are not statistically distinguishable across education groups. We also show that the results obtained using fixed education categories (e.g., high school graduates), rather than education quartiles, will be incorrect in important ways. Using fixed categories frequently misidentifies the groups with the worst mortality experiences and sometimes substantially overstates the increases in their death rates.

Our main finding is the characterizations of the most adverse mortality trends being concentrated among the least educated are overly simplistic and often inaccurate. One aspect of this is that the relationship between mortality trends and education varies sharply by gender. For women, the magnitudes of reductions in both logs and levels of death rates have generally increased monotonically with education. Conversely, males in the three lowest education quartiles often have fairly similar experiences. For both men and women, the average gap in mortality between the top 25 percent and the bottom 75 percent is growing.

These averages, nonetheless, conceal substantial disparities across groups. For example, we document the frequency with which the patterns are reversed within sex-age-race/ethnicity groups, such that lower education quartiles experienced larger mortality reductions than their more highly educated counterparts. There are also many cases where less-educated groups do worse but by amounts that are small enough that we fail to reject the null hypothesis of no difference, even when using a loose standard for rejection. These latter patterns may have been previously concealed because of the considerable attention paid to a small number of groups of low-educated non-Hispanic white (hereafter "white") males experiencing large increases in levels (but not logs) of death rates (Case and Deaton 2015, 2017; O'brien, Venkataramani, and Tsai 2017).

2 Conceptual Framework

Researchers commonly treat changes in health outcomes across population groups, in this case mortality patterns, as providing a useful indication of who is being most affected by health-related shocks. We examine this view, in the context of a model where mortality depends on investments in and depreciation of health capital (Grossman 1972). This section provides an overview, with further details in Appendix A. Consider an exogenous shock that produces an equal-size reduction in health capital for two groups varying in their initial stocks of capital. If the distribution of the baseline stock of health capital is monotonically increasing in education, or some other measure of socioeconomic status (SES), the *absolute* increase in death rates will be larger for the less educated group.[4] The intuition is that deaths are "left-tail" events, but more so for higher SES individuals, so that the relevant portion of the distribution over which deaths are being induced is "thicker" for low SES

[4] Our empirical specification uses educational attainment, rather than the more-encompassing "socioeconomic status". Our conceptual model, however, does not draw a distinction between education and SES.

groups. However, high SES individuals may, experience bigger *relative* increases in mortality risk from the same-sized negative shock.

These ideas are illustrated in Figure 1, which shows the left tail of cumulative distribution functions (CDF) for health capital, assumed here to be normally distributed and with the same variance but different means for high and low SES groups. The solid lines show the CDFs without a shock and the dotted lines are moved horizontally to the left by the shock, S, which is equal-sized for both groups. Death occurs if health capital falls below the threshold H^0, and the negative health shock increases the death rate by more for the low SES group ($R_l' - R_l$) than for high SES individuals ($R_h' - R_h$). However, since the no-shock death rates are so much lower for the latter group (R_h versus R_l), the relative increase is greater for them.

With more realistic assumptions, a uniform negative shock to some input (e.g. income) could reduce health capital more for low than high SES individuals, and so also increase the relative risk of death by a greater amount for them. For instance, this may occur because lower levels of overall health imply higher returns to given investments (or disinvestments) in health capital for low SES groups, or if they have less ability to mitigate the effects of the negative shock. Similarly, higher SES individuals might face lower prices for market inputs to health (e.g. through insurance), and so be more able to limit the reduction in health capital resulting from the shock.

The analysis of positive health shocks is largely the reverse of that just described, with the greatest absolute death rate reductions anticipated for the less educated and with unclear predictions for relative decreases. For instance, improvements in medical technology might result in larger total, but not necessarily percentage, mortality improvements for groups with higher baseline death rates.

This simple framework suggests that both absolute and relative changes in mortality can be informative for evaluating different questions. Changes in absolute death rates may be particularly relevant when attempting to target policy interventions to specific causes of death that have differential impacts across groups. Conversely, changes in relative rates may be more useful for understanding the underlying exposure and responses to health shocks. For this reason, we examine both absolute and relative changes in mortality.

3 Data and Methods

Our empirical analysis measures how mortality rates have changed over time for different population subgroups. This section first summarizes the data and methods used to construct mortality rates by education quartile, followed by a description of the regression specification.

3.1 Death Rates by Education Quartile

We calculate educational quartiles separately by 5-year age bins, gender, race/ethnicity and year. For compactness, we frequently abbreviate the first through fourth education quartiles as Q1 through Q4, where Q1 refers to the least-educated. This approach allows the distribution of education to differ across both demographic characteristics and time periods. We construct death rates for age group a, race/ethnicity r and education quartile i, in year t as:

$$mort_{arit} = \frac{deaths_{arit}}{pop_{arit}} \tag{1}$$

where $deaths_{arit}$ and pop_{arit} refer, respectively, to the number of deaths and population of the relevant group. We compute these death rates for five-year age groups ranging from 25 to 74.[5] In all cases, we calculate death rates separately for males and females, with the sex subscript excluded from equation (1) and later equations to simplify notation.

We use 2001-2017 data from the Centers for Disease Control and Prevention *Multiple Cause of Death* (*MCOD*) files to construct annual counts of deaths from all sources for specified age, education, sex and race-ethnicity groups.[6] Information on total population by age, gender, race/ethnicity, and year is obtained from the National Cancer Institute's *Surveillance Epidemiology and End Results* (*SEER*) database.[7] These are combined with information on the distribution of educational attainment by age, gender, race/ethnicity, and year using data from the Census Bureau's *American Community Survey* (*ACS*). We use 2001 as the starting year for our analysis because this is the first ACS year.

An empirical challenge in constructing death rates by education quartile is that education is measured in discrete units and the data are not fully comparable across time periods or sources. With continuous education measures, the exercise would be straightforward: we would simply calculate the group- and year-specific distribution of education and then divide it into quartiles. Instead, our approach is to use or construct single-year measures of education, ranging from 0 to 17 years.[8] Obtaining both deaths and population by single year of education and demographic group involves two complications. The first is that the *MCOD* and *ACS* files each record education in single year increments at various points throughout our sample period, but at other times code education into categories (e.g. high school, some college, college). We address this limitation using an

[5] The youngest group analyzed are 25-29 year olds because educational attainment is not meaningfully measured below this. Our oldest group are 70-74 year olds because (prior) mortality selection becomes an increasing issue at older ages.

[6] The *MCOD* files summarize information from the universe of death certificates to US residents.

[7] The SEER data are designed to supply more accurate population estimates for intercensal years than standard census projections.

[8] In two cases we combine groups. Persons with one year of college but no bachelor's degree are assigned as having 14 years of education. Those with one or more years of post-graduate education are categorized as having 17 years

imputation procedure, detailed in Appendix B, to estimate single year of education death and population counts. The second complication is that a single year of education may straddle quartiles. For example, if 22 percent of the group has 10 or fewer years of education, and 7 percent have exactly 11 years, the 25th percentile occurs somewhere between 10 and 11 years. Our procedure in these cases is to proportionately assign deaths from the overlapping education cell to each quartile, based on population shares. In the example just provided, three-sevenths of deaths and population for the 11-year group are assigned to Q1 and four-sevenths to Q2.[9]

An additional issue arising when computing results for a single race-ethnicity group is whether the education quartiles should be "general", constructed using the overall distribution of education for the specified sex-age group, or whether these should be "race-specific", calculated based only on educational attainment for individuals of the same race/ethnicity. Our primary analyses use general thresholds, reasoning that it does not make sense for two individuals with the same age, sex, and education to be placed in different quartiles because their races differ.[10] One implication is that the lower quartiles will be disproportionately populated by race/ethnicity groups with below average education levels (e.g. blacks and Hispanics) while the other groups (whites and other races) will be over-represented in the higher quartiles. For this reason, we also examine the robustness of our results to using race-specific education thresholds in Section 5.2.

[9] Three-sevenths are assigned to Q1 since $\frac{25-22}{29-22} = \frac{3}{7}$ and four-sevenths to Q2 since $\frac{29-25}{29-22} = \frac{4}{7}$.

[10] It seems reasonable to allow for differences in quartiles by age since, for example, the status of a 25-year-old high school graduate may be quite different than that of her 70-year-old counterpart.

3.2 Regression Specifications

Our analysis seeks to determine how mortality patterns have changed differentially over time across education quartiles. Towards this end, we first summarize changes across quartiles and then directly test for differential trends, by estimating:

$$mort_{arit} = \sum_{a=1}^{10} \sum_{r=1}^{4} \sum_{i=1}^{4} [\beta_{ari} group_{ari}] \ + \pi_1 trend + \sum_{i=2}^{4} [\ \pi_i trend \times Q_i] + \epsilon_{arit} \quad (2)$$

where $mort_{arit}$ is the level or log of the death rate for age group a, race/ethnicity r and education quartile i in year t; $group_{ari}$ is a group-specific fixed-effect, *trend* is a linear time trend, Q_2, Q_3, and Q_4 denote indicator variables for education quartiles 2, 3 and 4, and with the lowest quartile, Q_1, serving as the reference group. All of our regression models are estimated separately for men and women, so that each includes 160 age-race/ethnicity-education quartile groups. We cluster standard errors by age, race/ethnicity and education, and weight each cell by its population to obtain nationally-representative estimates.

In equation (2), π_1 captures the average mortality trend for the lowest education quartile (the reference group), which is restricted to be the same across all ages and races. The coefficient π_2 measures the mean differential change in trend for the second-lowest quartile relative to the lowest. The coefficients π_3 and π_4 have analogous interpretations for the third and fourth quartiles, relative to the first. The standard errors on $\hat{\pi}_2$ through $\hat{\pi}_4$ are used to calculate 95 percent confidence intervals on these differences relative to the lowest quartile.

Next, in order to examine how educational trends vary across specific age-race/ethnicity groups, we estimate:

$$mort_{arit} = \sum_{a=1}^{10} \sum_{r=1}^{4} \sum_{i=1}^{4} [\beta_{ari} group_{ari} + \delta_{ari} (group_{ari} \times trend)] + \ u_{arit} \qquad (3)$$

where all variables are as previously described. These regressions include separate time trends for all 160 age-race-/ethnicity education quartile groups (for each sex), without a constant or trend main effect, and so provide information on changes in death rates for each education quartile and group individually. The coefficients of primary interest, δ_{ari}, show the group-specific annual mortality changes, with β_{ari} indicating initial year levels.

4 Results

This section first describes overall trends in educational attainment and death rates by educational quartile for ages 25-74 combined. We construct aggregated measures of schooling and death rates in each year using common age standardization techniques described in Appendix B. Next, we summarize the overall pattern of education quartile differences in mortality trends, from 2001-2017, and then the frequency with which group-specific education trends are statistically distinguishable from each other. Section 5 follows with extensions and robustness tests.

4.1 Aggregate Trends

Mean educational attainment across all ages for Q1 varied from 9.7 to 10.2 years, depending upon the year, while the corresponding averages for Q2, Q3 and Q4 were 12.1 to 12.5, 13.9 to 14.6, and 16.3 to 16.5 years respectively (Appendix Figure C1). Schooling rose modestly over time for men and somewhat more for women: average male education grew by 0.46, 0.26, 0.52 and 0.07 years from 2001-2017 for Q1 through Q4, compared to 0.46, 0.63, 0.90 and 0.35 years for females. These overall changes sometimes conceal much larger growth. For example, average educational attainment increased by 1.3 and 1.0 years respectively for 60-64 year old Q1 and Q3 men and by 1.2, 1.6 and 0.8 years for same-aged Q1, Q3 and Q4 women.

Overall sex-specific death rates from 2001-2017 for the four education quartiles are shown in Figure 2. The gradient between death rates and educational attainment is clear, but with higher absolute rates for men. For both men and women, Q4 is the only quartile to experience steady declines since 2001. Death rates fell for Q3 males and females until 2010, and then subsequently rose. Trends for the bottom two quartiles were flat or slightly increasing until 2010, and then exhibited a sharper increase.

4.2 Econometric Estimates

We begin our econometric estimation by examining the overall pattern of trends in mortality rates across education quartiles. Figure 3 presents results from equation (2), which aggregates the quartile trends across groups. The entry for Q1 is the trend "main effect", $\hat{\pi}_1$, from the model. The point estimates for the remaining groups are calculated as the main effect just described, plus the quartile-specific trend coefficients. For instance, the estimated trend for Q2 is calculated as $\hat{\pi}_1 + \hat{\pi}_2$. The 95 percent confidence intervals (CIs) are centered on the Q2 through Q4 total effects and indicate whether the corresponding trend is statistically significantly different from the Q1 trend. (This can be visually observed by whether the CI crosses the dotted horizontal line showing the estimated Q1 trend).

For men, trends in both logs and levels of death rates are flat for the lowest quartile (Figures 3a and 3c) and the differences between Q1, Q2 and Q3 are never statistically significant. Conversely, Q4 experiences reductions in log death rates that are significantly larger than for Q1 and decreases in death rates that are marginally greater. Among women, there is more evidence of monotonic trends across quartiles. Log death rates for the lowest quartile rose, on average, by nearly 1 percent (Figure 3d). Each higher quartile had successively better mortality trends. For instance, death rates in the top two quartiles fell by about 6 and 8 per 100,000 and logs and levels of death rates improved by statistically significantly larger amounts for Q3 and Q4, than for Q1.

The specifications just estimated are restrictive, since they treat the education trends as being constant across all groups. To allow for more flexible trends, we next provide a series of plots visually summarizing how the distribution of estimated trends differs by education. These distributions, displayed in Figure 4, present the trend coefficients from estimating equation (3), ranked from largest increase to greatest decrease within each education quartile. Thus, each point represents the estimate of δ_{ari} for the particular education group, with a total of 40 estimates for each quartile (10 ages x 4 races).

The figure provides two main results. First, there is a clear overall education gradient in mortality trends for women: the biggest reductions in log death rates are observed for Q4, followed by Q3 and with the least favorable changes for the two lowest quartiles. For men, by contrast, while Q4 experienced the largest average mortality declines, there is often little difference between Q1, Q2, or Q3.[11]

Beyond these general patterns, there is often substantial overlap of the estimated trends across the education quartiles. To illustrate, observe that the fifty percent of the 160 groups (for each sex) with estimates between the 25th and 75th percentile of estimated mortality trends are fairly evenly distributed across education quartiles. For death rates, 17, 18, 18 and 27 groups of Q1, Q2, Q3 and Q4 males are within the interquartile range as are 19, 15, 19 and 27 groups of corresponding females. For log death rates, the corresponding numbers are 20, 23, 19 and 18 for males and 17, 20, 25 and 18 for females. Outside of the interquartile range, the greatest mortality declines are concentrated among Q4 and the highest increases among the two lowest quartiles (see Figure 4). Nonetheless, the results demonstrate similar mortality trends across education quartiles for many groups in the middle of the distribution.

[11] With that said, the three groups with the largest increases are whites in the lowest education quartile aged 30-34, 50-54, and 55-59.

4.3 Testing for Worse Mortality Trends Among the Lower Education Quartiles

The results above indicate aggregate patterns across education groups but are somewhat limited in that they compare trends across all age and race-ethnicity groups or for the overall distribution. However, it is possible, for example, that mortality trends are worse in the lower education quartiles for some of these groups but not others in ways not identified by the analysis performed up to this point. With this in mind, we next examine the heterogeneity in results across education quartiles within groups. For clarity, we define the mortality trends to be *non-monotonic* if an education quartile has a slower growth or larger decline in mortality than a higher quartile for the same age, race, and sex. Thus a non-monotonicity occurs if: Q1 does better than Q2, Q3 or Q4; Q2 does better than Q3 or Q4; or Q3 does better than Q4. We also evaluate whether these differences are statistically distinguishable by using the estimated group-specific trend coefficients to compare: Q1 vs. Q2, Q3 and Q4; Q2 vs. Q3 and Q4; and Q3 vs. Q4, with the standard errors and associated p-values estimated through bootstrapping, taking 10,000 samples with replacement to calculate a distribution of trend estimates. Here, we conduct 1-sided tests where the null hypothesis is that less-educated groups have equal or better mortality trends than the higher quartiles and the alternative hypothesis is that they have worse trends. The choice of 1-sided tests is guided by the prevailing narrative and allows us more power to reject the null when a lower-educated quartile has fared worse than a higher quartile.

A simple comparison of the trend coefficients reveals that non-monotonicities occur for 30 of 40 groups of males for trends in log death rates for 16 of 40 groups of females. For death rates, non-monotonicities occur for 31 and 25 groups of males and females, respectively. When examining log death rates, the non-monotonicities most frequently occur because Q3 does worse than either Q1 or Q2, while both Q3 and Q4 are often observed to have worse trends than Q2 for death rates. Additional details are provided in Appendix Table C1.

We next examine the frequency with which the mortality trends of lower quartiles are statistically distinguishable from those of higher quartiles, based on the 1-sided tests described above. Figure 5 summarizes the results of these tests by plotting the CDF of the *p*-values of the hypothesis tests, separately by sex and by logs versus levels for the six possible test for each group (Q1 vs. Q2-Q4, Q2 vs. Q3-Q4 and Q3 vs. Q4). The dashed 45-degree line represents the distribution of *p*-values under the "grand null" that lower-educated quartiles have experienced trends no worse than those of higher-educated quartiles of the same group, using data for each of the six tests just described. By definition, with trends that were randomly distributed across education quartiles, 5 percent of tests would have *p*-values less than 0.05, 10 percent below 0.1, and so forth. Also, note that the *p*-values will be above 0.5 in cases where the lower quartile has a better estimated trend than the higher quartile, since the test is 1-sided.

As expected, given the large mortality declines of Q4 observed above, the evidence against the grand null hypothesis of equal trends across educational quartiles is strong. For instance, for log death rates, 48 percent of male and 65 percent of female tests have *p*-values below 0.05. For death rates, 33 percent of male and half of female tests have *p*-values below 0.05.

At the same time, evidence in support of a widening gradient throughout all quartiles is often weak. This is especially true for men. In 26 percent of tests for log death rates and 42 percent for death rates, the *p*-value exceeds 0.5, indicating that the estimated trends for lower-educated quartiles are more favorable than for higher-educated quartiles in the same group. For females, 15 and 30 percent of the tests have *p*-values over 0.5 for death rates and log death rates, respectively. Moreover, 20 percent of tests for male death rates have a *p*-value *above* 0.98 and 17 percent of tests for female death rates have *p*-values above 0.9.

These disproportionate shares of p-values close to 1 are displayed by the CDFs in Figure 5 crossing the 45-degree line at the upper centiles of the distribution.

Moreover, we often fail to detect statistically distinguishable differences when allowing for less stringent criteria for rejecting the null hypothesis. For instance, when using a p-value of 0.2—well above conventional levels of statistical significance—the null is not rejected in 39, 58, 26 and 42 percent of cases for male logs, male levels, female logs and female levels of mortality rates. It is important to note that we have more statistical power when considering log death rates than death rates, which explains why we fail to reject more often in the latter.[12]

Since the top quartile generally has the greatest declines in mortality, we next test for differences between the bottom two quartiles and third quartile by considering two tests for each group: Q1 vs. Q3 and Q2 vs. Q3. The results are summarized in Figure 6. For men, p-values exceed 0.5 over half the time, for both logs and levels, indicating the frequency with which Q3 does worse than one of the lower education quartiles, while this occurs 19 and 30 percent of the time for females. Using the p-value of 0.2 as the threshold for hypothesis testing, we are able to reject the null hypothesis of greater declines for less-educated quartiles just 35 and 36 percent of the time for male logs and levels of death rates, compared to in 71 and 64 percent of cases for females.

The fourth quartile frequently has experienced better trends than Q1 and Q2, as shown in Appendix Figure C2. For log death rates, the pattern is nearly universal. Lower

[12] As an example, we simulate power to detect a 1 percent change in mortality between two quartiles for a group with a baseline death rate of 750 per 100,000. Assuming a normally distributed random error equal to 10 percent of the baseline mean and a type I error rate of 0.05, we have 87.2% power to detect a change of 0.01 for log death rates but only 42.6% power to detect a change of 7.5 for death rates. For a larger type I error rate of 0.2, we have 97.7% power for log death rates and 72.3% power for death rates.

quartiles, however, have larger reductions in death rates in about a third of cases for both men and women. This pattern can be explained by the low baseline rates for Q4.

5 Robustness and Extensions

This section extends upon the preceding results in several ways. First, we assess whether the observed patterns are concealed or incorrectly estimated when using fixed categories of educational attainment, rather than education quartiles. Next, we replicate our estimates using race-specific rather than general thresholds for classifying education quartiles. Third, we investigate how sampling variation in the ACS may affect our estimates. Fourth, we split the lowest education quartile into those below and above the bottom decile to examine a different classification of the lowest-educated group. Finally, we present descriptive evidence examining whether geography may explain our findings.

5.1 Education Quartiles vs. Categories

It is considerably more difficult to analyze education quartiles – whose thresholds vary across groups and over time – than to use fixed education categories (e.g. high school graduate). Here we summarize whether such extra effort is warranted by examining whether education quartiles produce meaningfully different results compared to fixed categories. Appendix D provides detail on this analysis and the findings.

The correlations between the quartile and categorical mortality trend coefficients are reasonably high, between 0.72 and 0.90, and the highest group (college graduates) continues to experience the largest mortality declines, as others have documented (e.g. Case and Deaton 2017). Yet mortality in the second-highest category (some college) is substantially worse compared to the quartile-based results. In fact, log death rates for males with some college have risen by more than for those never attending college (Appendix Figure D1). These comparisons by category are confounded by compositional changes, unlike our quartile-based estimates.

Important differences also emerge when considering mortality trends for individual groups. The magnitudes of the trend increases for those with the worst experiences are often dramatically overstated when using education categories. For instance, less-educated whites near retirement age experienced the largest growth in death rates with either classification method, but the use of educational categories overstates the increases by a factor of two or three (Appendix Table D1). In the most extreme cases, the estimated rise in death rates using educational categories are an order of magnitude larger than those based on quartiles. Our overall conclusion is that the gains from classifying groups by education quartile are substantial and justify the greater complexity involved.

5.2 Race-specific education thresholds

Our main specifications used age- and sex- but not race-specific thresholds for categorizing education quartiles. Appendix Figures C3 and C4 present quantile plots for logs and levels of death rates, allowing the quartile thresholds to also differ by race. The monotonic pattern between education quartiles and mortality trends for women is somewhat weakened, with almost half of Q1 groups now overlapping with Q2. However, with one exception, the largest increases continue to be observed in the lowest education quartile. For men, there continue to be similar patterns between Q1 and Q2, but with Q3 log death rates now closer to those of Q4. For both levels and logs, non-monontonic patterns remain common for men.

5.3 Sampling variation in education shares

We use survey data from the ACS to estimate group-specific education shares. Although the ACS is designed to be nationally representative with weighting, small cell sizes thus could lead to substantial sampling error in these shares and, consequently, in the denominator of our mortality rate calculations. We investigate this possibility by calculating the standard errors of the shares for four education categories.[13] Appendix Figure C5 displays histograms of the coefficient of variation (CV) of these shares by sex-age-race/ethnicity group. Most are estimated with a high degree of precision: the median CV is 33 in 2001 and 64 in 2017. The 5th percentiles are 13 in 2001 and 31 in 2017. The small CV's provide support to our main results, and indicate that sampling variation in education shares is unlikely to change their interpretation.

5.4 Decomposing Mortality Experiences of the Bottom Quartile

Novosad, Rafkin and Asher (2020), hereafter NRA, have recently argued that the worst mortality experiences have been concentrated among the lowest decile of whites and blacks. To compare our results, we estimate models where the bottom quartile is split into those at or below and versus above the 10th percentile. Appendix Figures C6 and C7 show that the bottom 10 percent of non-whites almost always fared *better* than the 11th to 25th percentiles. Conversely, for whites the lowest decile almost always had worse experiences than their slightly more educated counterparts for death rates, consistent with NRA, but with mixed results obtained when examining log death rates.

Differences between our findings and those of NRA for non-whites and may partially reflect the data sources used to calculate group-specific populations. NRA utilize the Current Population Survey (CPS), whereas we obtain overall sex-age-race populations from the (more accurate) SEER data and then use data from the ACS to estimate education shares for each group. Neither the CPS nor the ACS is fully representative at this level, but the ACS has a substantially larger sample size.[14] Small sample sizes may lead to instability

[13] Although our analysis uses single years of education to construct quartiles, we calculate the CV for four education categories here since our interest is in potential mismeasurement of the cutoffs between quartiles, rather than between each particular year of schooling.

[14] For example, the CPS collects data from 100,000 residences compared to 3 million in recent years of the ACS. See https://www.census.gov/topics/income-poverty/poverty/guidance/data-sources/acs-vs-cps.html for more information comparing the two surveys.

that produces inaccurate estimates of the changes over time in population and thus mortality rates. Also, the sampling weights are based on age, race and sex but *not* education, introducing potential problems when using them to construct education-specific populations. The CPS further excludes those who are institutionalized and military persons living in group quarters, requiring NRA to make adjustments for this population.

We explore these issues by calculating, for each group, the ratio of the population estimates obtained from the CPS alone versus those from the combined SEER and ACS data. Appendix Figure C8 plots these ratios separately for whites and blacks in 2001 and 2017. A ratio of 1.2, for example, means the CPS population estimate is 20 percent larger than that from the SEER/ACS. The ratios change substantially over time for blacks, and less so for whites, which may explain why NRA obtain different findings for blacks than we do for non-whites.[15] Since NRA focus exclusively on whites and blacks, we also recalculate our quartile thresholds after excluding Hispanics and other races. As shown in Appendix Figure C9, we find broadly similar patterns to our main results in Figure 4, except that the male log death rate for the third quartile is generally higher than for Q1 or Q2.

Disparate results for the bottom decile of non-whites may also reflect the time periods analyzed. NRA study mortality changes between 1992 and 2015, while we focus on 2001 to 2017. Death rates for prime-age blacks declined substantially towards the end of the 20th centure, part because of better treatments for HIV (Coile and Duggan 2019), before our analysis period. Conversely, deaths due involving fentanyl increased extremely rapidly towards the end of the timespan we analyze (Agency for Health Care Research and Quality 2020), particularly in 2016 and 2017 which are not included in NRA's estimates.

[15] The dissimilar qualitative results for low-educated blacks could also be due to NR's earlier (1992) starting date or because of the bounding procedure they employ, which differs from our method of proportional assignment. We discuss the trade-offs in these modeling decisions in Appendix B.

5.5 Geographic Patterns

Recent literature suggests a possible role for geographic differences in explaining the mortality growth of the less educated (Montez et al. 2019; Montez, Hayward, and Zajacova 2019). However, the analysis of regional patterns is complicated since migration is endogenous, introducing another potential source of selection bias (Currie and Schwandt 2016). To explore this issue, we calculate the mix of education quartiles within Census divisions over time and determine that there is generally little change in the quartile shares from 2001 to 2017 (Appendix Figure C10).

Next, we draw on Woolf and Schoomaker's (2019) calculations of relative changes in mortality across states. The data reveal mixed evidence on a Census division's relative changes in death rates and its level of educational attainment (Appendix Figure C11). Specifically, census divisions with larger Q4 shares had greater percentage reductions in death rates, while the reverse is true for areas with larger population shares of Q2. Both results could, in principle, be consistent with some correlated geographic factors providing the actual source of mortality differences. However, the relationship is flat for Q1, rather than positive as it is for Q2, which operates against such a story. Taken together, this analysis provides limited support for geography-based explanations for our main results.

6 Discussion

We provide a detailed analysis of mortality changes, from 2001-2017, for subgroups stratified by sex, race-ethnicity, 5-year age groups and education quartiles. Our most important conclusion is that prior characterizations suggesting that the worst mortality trends have been concentrated in the bottom of the education distribution are overly simplified and, in important ways, substantially incorrect. Consistent with prior work, we find the least-educated women generally experienced the smallest mortality reductions, and often sizable increases. However, this is less true for males, where all but the highest education quartile had substantially similar patterns. Conversely, both men and women in the top quartile, on average, had the most favorable mortality experiences, leading to a widening gap between those at the top and the rest of the population. The second key point is that there is substantial variation around these averages. Within sex-race/ethnicity-age groups, monotonicity violations are frequent—whereby lower education quartiles have more favorable mortality experiences than higher ones—and even when the less educated do worse, the differences are often small and statistically indistinguishable from those of the higher quartiles.

Our results should be interpreted in light of several limitations. First, causation is always difficult to infer from descriptive analyses, although an understanding of the "stylized facts" provides an important first step. Second, data limitations restrict the time period analyzed and it would be informative to test the sensitivity of the results to the use of different starting and ending years (Coile and Duggan 2019). Third, our focus is on estimating linear trends to concisely summarize mortality patterns across groups, although non-linearities could exist for some groups. We have also not accounted for immigration, which could be relevant if mortality rates differ between foreign-born and native-born populations. However, since net migration has been largest for Hispanics over the analysis period and the main results are qualitatively similar when we only consider whites and blacks (although with somewhat worse experience for Q3 males), immigration patterns are probably not driving our findings. Finally, our methods are intended to account for rising levels of education, but we have not considered the causal role that education itself may play on mortality. While conceptually important, we observe the largest increases in educational attainment over time for the lower quartiles, but worse average mortality trends for them. This suggests that within-quartile changes in educational attainment are unlikely to meaningfully affect the interpretation of our results.

Throughout the analysis, we stress the utility of considering both relative and absolute changes in death rates and believe that prior research has probably paid too little attention to the former. In particular, since the absolute growth or decline equals the baseline level times the percentage change, groups with the largest absolute changes will often be those with high initial death rates. For instance, some of the biggest *percentage* mortality reductions occurred among 35-54 year-old men, while the *absolute* decreases were substantially smaller for them than for corresponding 65-74 year olds. However, this primarily results from the strong positive age gradient in death rates and it would almost certainly be incorrect to conclude that progress in reducing mortality was smaller for the former group than for the latter.

Our conceptual framework demonstrates that equal-sized shocks to health capital are likely to result in the largest absolute changes in death rates for lower educated groups, while the differences in percentage changes are likely to be more muted or even reversed. Our empirical findings are generally *not* in accordance with these predictions. Specifically, mortality has trended downwards for the majority of groups over time, suggesting positive shocks to health capital, but the absolute reductions are often larger for the most highly educated. This pattern suggests differential shocks, or responses to these shocks, by education level. Two other pieces of evidence support this possibility. First, while most mortality rates have fallen over time, on average, increases were observed for over one-fifth of the groups examined.[16] Second, as highlighted above, the results for education quartiles within age-race/ethnicity-sex groups are often quite heterogeneous.

A natural question is what factors explain these particular mortality patterns. The relatively unfavorable mortality trends of the second and third education quartiles for men

[16] Across both sexes, there is only one group out of 80, 30-34 year old white men, for whom death rates increased for each quartile (ranging from 7.0 deaths per 100,000 for Q1 to 0.2 per 100,000 for Q4).

are initially surprising, but may be consistent with recent evidence indicating poor outcomes of persons with some college experience but who did not graduate (Zajacova, Rogers, and Johnson-Lawrence 2012; Zajacova and Lawrence 2018). Although the reasons for this are not well understood, one possibility is that individuals completing some college without receiving four-year degrees are relatively poor and come from disadvantaged backgrounds. Cherlin (2018) emphasizes the relative decline in job opportunities and eroding stability of families among the moderately educated. These patterns have the potential to differentially affect males and females. In future work, we intend to examine specific causes of death and patterns by age and race that may shed additional light on the trends observed here.

References

Agency for Health Care Research and Quality. 2020. "Blacks Experiencing Fast-Rising Rates of Overdose Deaths Involving Synthetic Opioids Other than Methadone." AHRQ Publication No. 20-0032-1.

Autor, David, David Dorn, and Gordon Hanson. 2019. "When Work Disappears: Manufacturing Decline and the Falling Marriage-Market Value of Young Men." *American Economic Review: Insights* 1 (2): 161–78.

Bosworth, Barry. 2018. "Increasing Disparities in Mortality by Socioeconomic Status." *Annual Review of Public Health* 39: 237–51.

Bound, John, Arline T. Geronimus, Javier M. Rodriguez, and Timothy A. Waidmann. 2015. "Measuring Recent Apparent Declines in Longevity: The Role of Increasing Educational Attainment." *Health Affairs* 34 (12): 2167–73. https://doi.org/10.1377/hlthaff.2015.0481.

Case, Anne, and Angus Deaton. 2015. "Rising Morbidity and Mortality in Midlife among White Non-Hispanic Americans in the 21st Century." *Proceedings of the National Academy of Sciences* 112 (49): 201518393. https://doi.org/10.1073/pnas.1518393112.

———. 2017. "Mortality and Morbidity in the 21st Century." *Brookings Papers on Economic Activity.*

Cherlin, Andrew J. 2018. "How Inequality Drives Family Formation The Prima Facie Case." In *Unequal Family Lives: Causes and Consequences in Europe and the Americas*, edited by Naomi R Cahn, June Carbone, Laurie Fields DeRose, and J Bradford Wilcox, 69–82. Cambridge: Cambridge University Press.

Clark, Damon, and Heather Royer. 2013. "The Effect of Education on Adult Mortality and Health: Evidence from Britain." *American Economic Review* 103 (6): 2087–2120.

Coile, Courtney C, and Mark G Duggan. 2019. "When Labor's Lost: Health, Family Life, Incarceration, and Education in a Time of Declining Economic Opportunity for Low-Skilled Men." *Journal of Economic Perspectives* 33 (2): 191–210.

Currie, Janet, and Hannes Schwandt. 2016. "Mortality Inequality: The Good News from a County-Level Approach." *Journal of Economic Perspectives* 30 (2): 29–52.

Cutler, David M., and Adriana Lleras-Muney. 2010. "Understanding Differences in Health

Behaviors by Education.” *Journal of Health Economics* 29 (1): 1–28. https://doi.org/10.1016/j.jhealeco.2009.10.003.

Cutler, David M, Fabian Lange, Ellen Meara, Seth Richards-Shubik, and Christopher J Ruhm. 2011. “Rising Educational Gradients in Mortality: The Role of Behavioral Risk Factors.” *Journal of Health Economics* 30 (6): 1174–87.

de Walque, Damien. 2010. "Education, Information, and Smoking Decisions: Evidence from Smoking Histories in the United States, 1940-2000. *Journal of Human Resources* 45(3): 682-717.

Dowd, Jennifer B, and Amar Hamoudi. 2014. “Is Life Expectancy Really Falling for Groups of Low Socio-Economic Status? Lagged Selection Bias and Artefactual Trends in Mortality.” *International Journal of Epidemiology* 43 (4): 983–88. https://doi.org/https://doi.org/10.1093/ije/dyu120.

Goldring, Thomas, Fabian Lange, and Seth Richards-Shubik. 2016. “Testing for Changes in the SES-Mortality Gradient When the Distribution of Education Changes Too.” *Journal of Health Economics* 46: 120–30. https://doi.org/10.1016/j.jhealeco.2015.12.002.

Grossman, Michael. 1972. “On the Concept of Health Capital and the Demand for Health.” *Journal of Political Economy*. https://doi.org/10.1086/259880.

Kitagawa, Evelyn M, and Philip M Hauser. 1973. *Differential Mortality in the United States: A Study in Socioeconomic Epidemiology*. Cambridge, MA: Harvard University Press.

Lleras-Muney, Adriana. 2005. “The Relationship between Education and Adult Mortality in the United States.” *The Review of Economic Studies* 72 (1): 189–221.

Marmot, Michael G, Martin J Shipley, and Geoffrey Rose. 1984. “Inequalities in Death—Specific Explanations of a General Pattern?” *The Lancet* 323 (8384): 1003–6.

Marmot, Michael G, Stephen Stansfeld, Chandra Patel, Fiona North, Jenny Head, Ian White, Eric Brunner, Amanda Feeney, and G Davey Smith. 1991. “Health Inequalities among British Civil Servants: The Whitehall II Study.” *The Lancet* 337 (8754): 1387–93.

Meara, Ellen R, Seth Richards, and David M Cutler. 2008. “The Gap Gets Bigger: Changes in Mortality and Life Expectancy, by Education, 1981–2000.” *Health Affairs*

27 (2): 350–60.

Montez, Jennifer Karas, Mark D Hayward, and Anna Zajacova. 2019. "Educational Disparities in Adult Health: US States as Institutional Actors on the Association." *Socius* 5: 1–14.

Montez, Jennifer Karas, Robert A Hummer, Mark D Hayward, Hyeyoung Woo, and Richard G Rogers. 2011. "Trends in the Educational Gradient of US Adult Mortality from 1986 through 2006 by Race, Gender, and Age Group." *Research on Aging* 33 (2): 145–71.

Montez, Jennifer Karas, Anna Zajacova, Mark D Hayward, Steven H Woolf, Derek Chapman, and Jason Beckfield. 2019. "Educational Disparities in Adult Mortality Across US States: How Do They Differ, and Have They Changed Since the Mid-1980s?" *Demography*, 1–24.

Murphy, Sherry L., Jiaquan Xu, Kenneth D Kochanek, and Elizabeth Arias. 2018. "Mortality in the United States, 2017." *NCHS Data Brief, No. 328.*

National Center for Health Statistics. 2017. "Health United States, 2017: With Special Feature on Mortality." Hyattsville, MD.

Novosad, Paul, Charlie Rafkin, and Sam Asher. 2020. "Mortality Changes Among Less Educated Americans." Mimeograph, Darmouth University.

O'brien, Rourke L, Atheendar S Venkataramani, and Alexander C Tsai. 2017. "Economic Mobility and the Mortality Crisis Among US Middle-Aged Whites." *Epidemiology (Cambridge, Mass.)* 28 (2): e12.

Olshansky, S. Jay, Toni Antonucci, Lisa Berkman, Robert H. Binstock, Axel Boersch-Supan, John T. Cacioppo, Bruce A. Carnes, et al. 2012. "Differences in Life Expectancy Due to Race and Educational Differences Are Widening, and Many May Not Catch Up." *Health Affairs* 31 (8): 1803–13. https://doi.org/10.1377/hlthaff.2011.0746.

Preston, Samuel H, and Haidong Wang. 2006. "Sex Mortality Differences in the United States: The Role of Cohort Smoking Patterns." *Demography* 43 (4): 631–46.

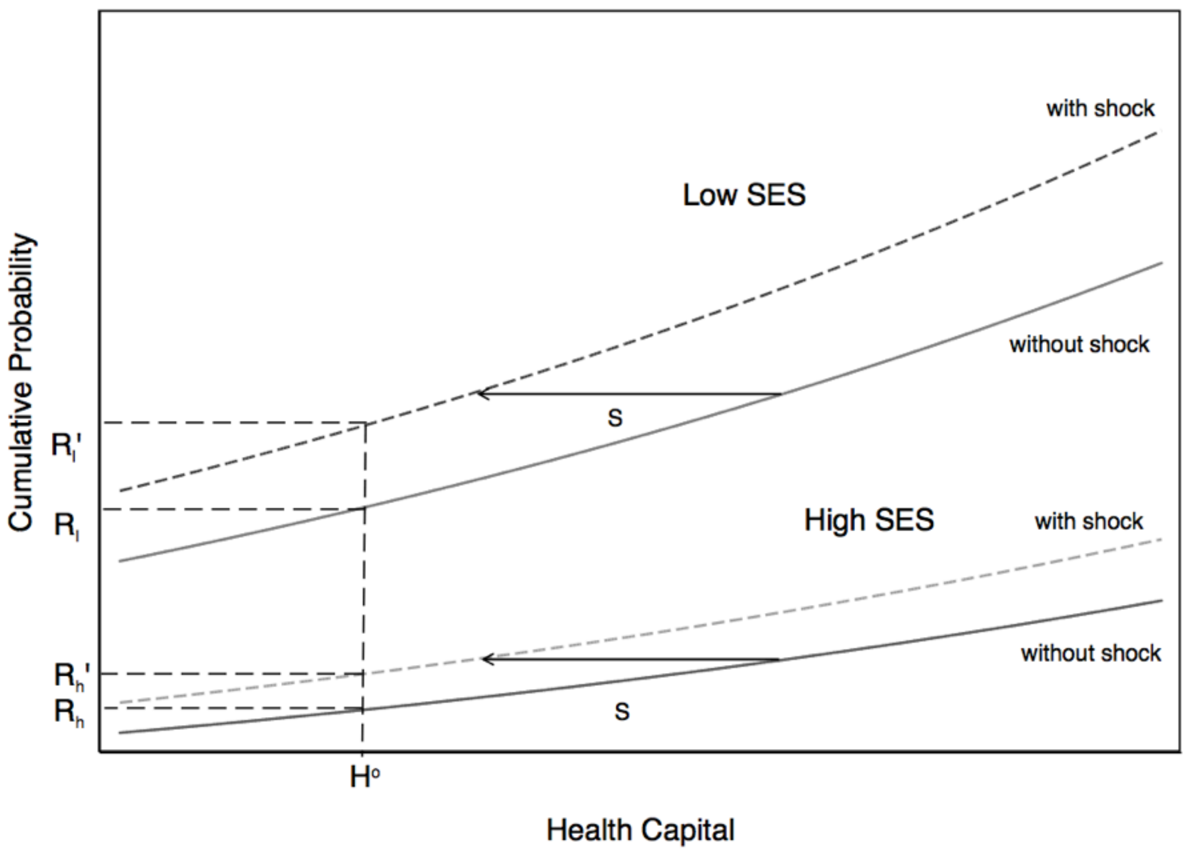
Rostron, Brian L, Elizabeth Arias, and John L Boies. 2010. "Education Reporting and Classification on Death Certificates in the United States." *Vital Health and Statistics* 2 (151). www.cdc.gov/nchs/data/series/sr_02/sr02_151.pdf.

Woolf, Steven H., and Heidi Schoomaker. 2019. "Life expectancy and mortality rates in the United States, 1959-2017." *JAMA* 322(20): 1996-2016.

Zajacova, Anna, and Elizabeth M Lawrence. 2018. "The Relationship between Education and Health: Reducing Disparities through a Contextual Approach." *Annual Review of Public Health* 39: 273–89.

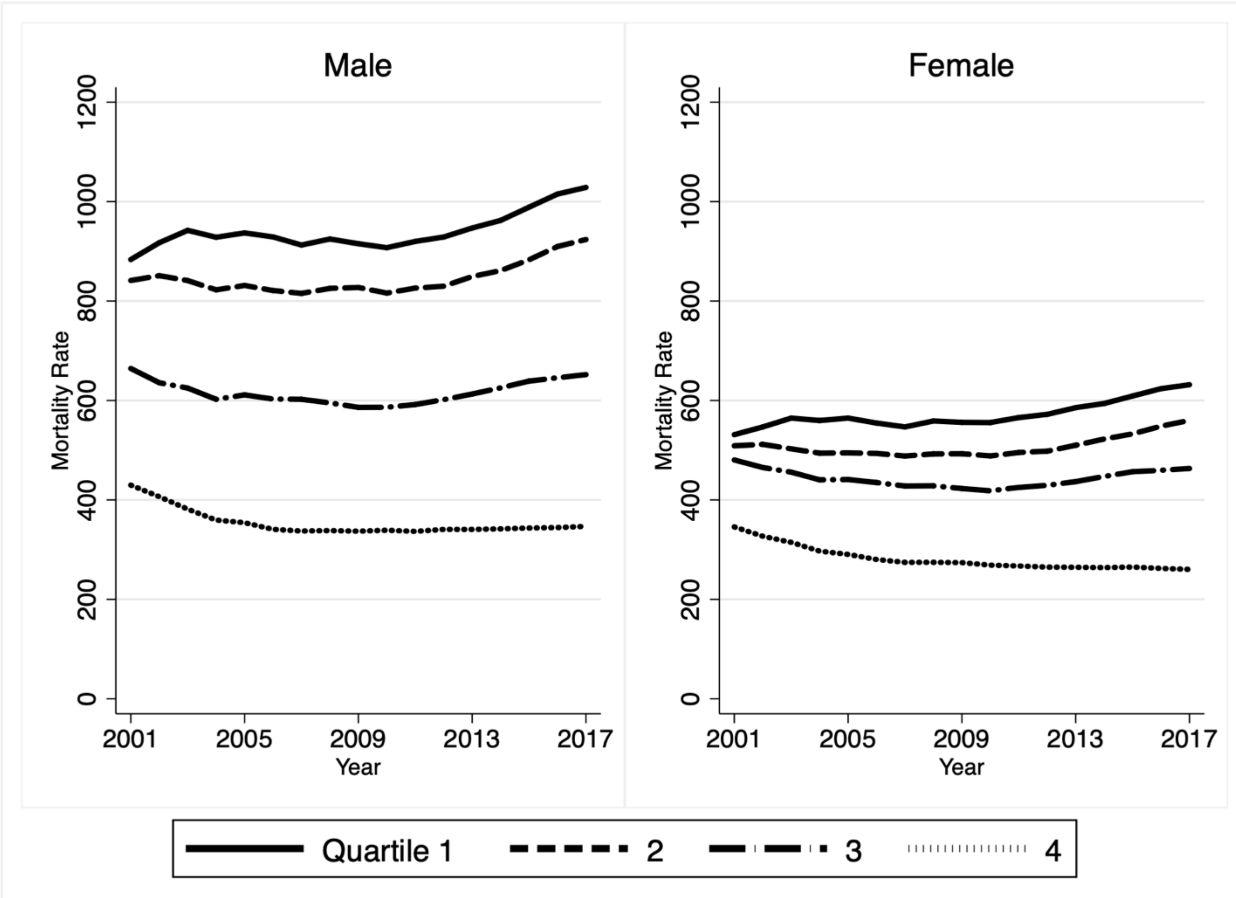
Zajacova, Anna, Richard G Rogers, and Vicki Johnson-Lawrence. 2012. "Glitch in the Gradient: Additional Education Does Not Uniformly Equal Better Health." *Social Science & Medicine* 75 (11): 2007–12.

Figure 1. Probability of death with and without health shock by SES



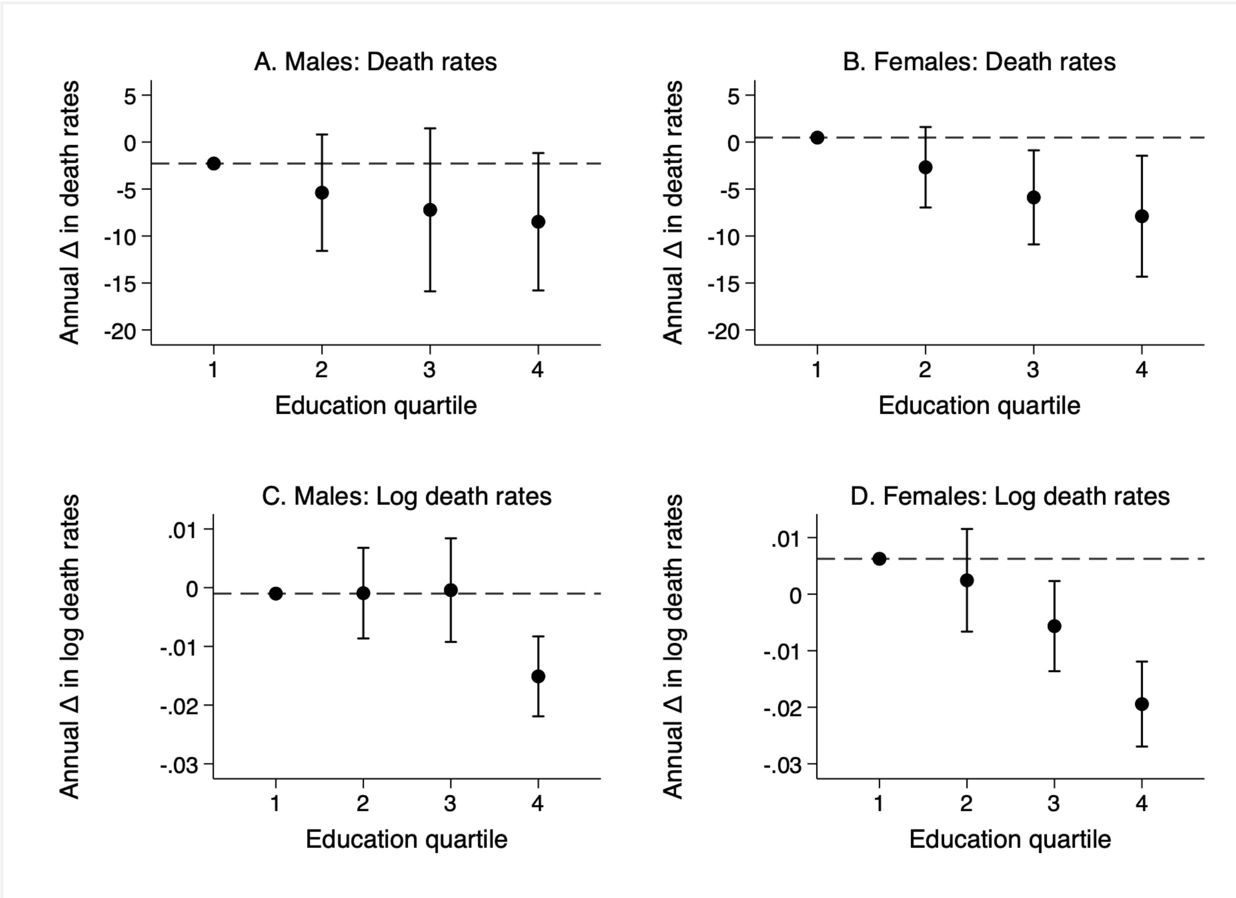
Note: Figure shows the cumulative distribution function for health capital which is assumed to have an equal variance for high and low SES groups but with a higher mean value for the former group. A health shock S causes an equal left-ward shift of the distribution for both groups. Death occurs if health capital falls below H^0, and the fraction of the group dying is shown by the dotted lines extending to the y-axis. Figure shows the case where normal distributions in the "without shock" case have means of 0 and 1, equal variances of 1.0 and where the health shock shifts both distributions to the left by 0.075 standard deviations.

Figure 2. Death rates by education quartile and sex, 25-74 year olds



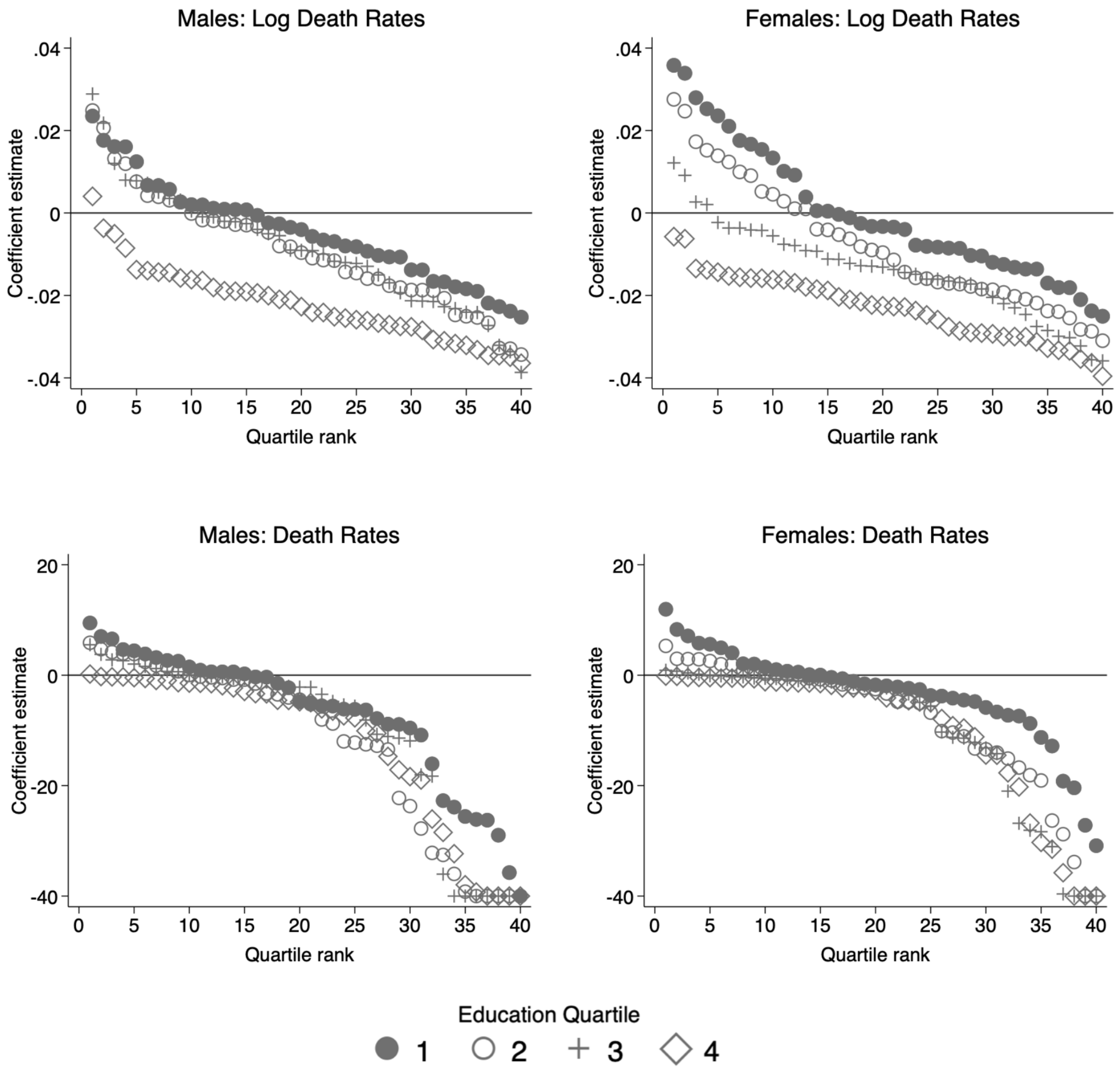
Note: Figure shows age-adjusted death rates per 100,000 by education quartile among 25-74 olds from 2001-2017. Death rates are calculated using MCOD, ACS, and SEER data as described in Appendix B. Death rates are age-standardized using the 2017 age distribution specific to each education quartile. Education quartiles are calculated separately by sex, age and year.

Figure 3. Estimated Trend Differences by Quartile



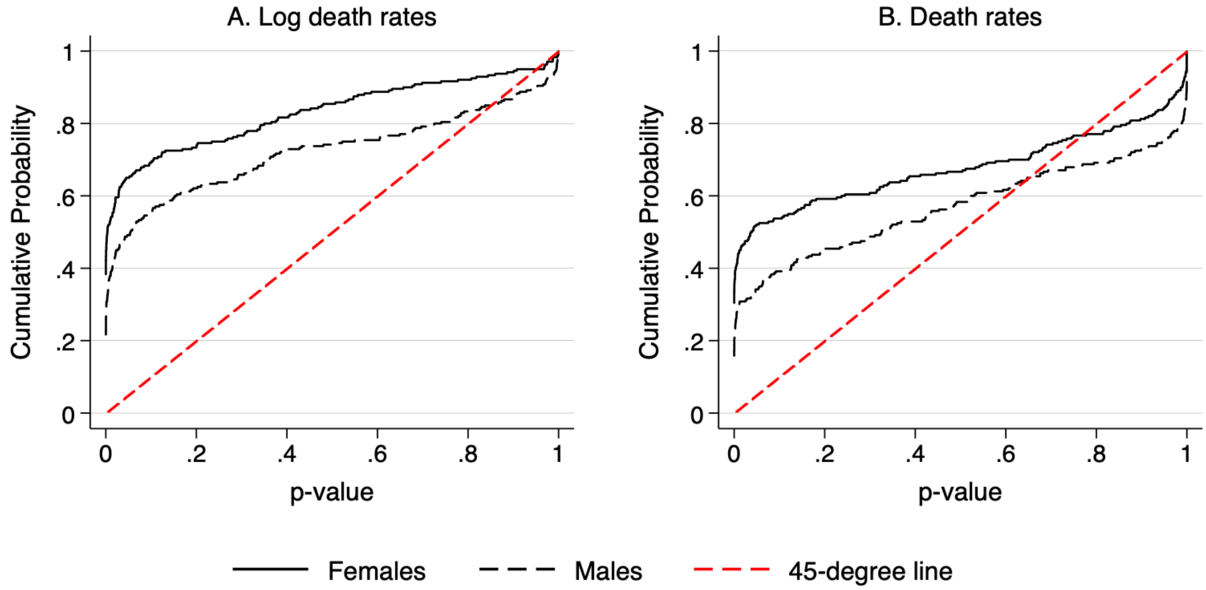
Note: Figure shows regression results from estimating equation (2). Dots represent estimates on the trend coefficient for each quartile, with those for quartiles 2, 3 and 4 calculated by adding the estimate on the corresponding regression trend interaction term to the trend estimate corresponding to quartile 1. The whiskers plot the 95% confidence interval of the difference relative to quartile 1, with the difference centered on the mean for quartiles 2, 3, and 4. If the upper-bound of the confidence interval overlaps with the horizontal dotted line, then the quartile's trend is not statistically distinguishable from the trend for quartile 1. Regression is weighted using the population in each group (age/sex/race/quartile cell), and standard errors are clustered at the group level.

Figure 4. Ranked mortality trends by education quartile



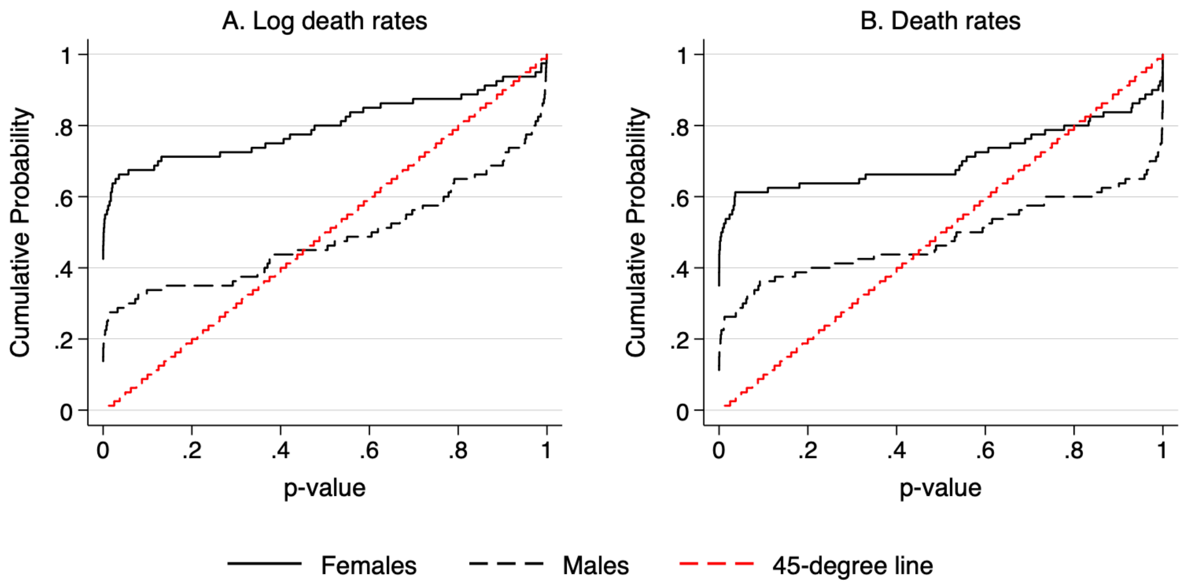
Note: Figure shows quantile plots of the trend coefficient estimates on log and levels of death rates from estimating equation (3). Coefficient estimates are reported separately by education quartile, pooling races and ages together. The x-axis lists the rank of the trend coefficient estimate within each education quartile, in which higher ranks denote larger reductions in mortality rates. For clarity, we bottom-code decreases in death rates at negative 40 to permit visible detection of differences for most of the distribution. These decreases are presented in Appendix Table D.3.

Figure 5. CDFs of p-values of tests of non-monotonicity in mortality trends



Note: Figure plots the cumulative distribution functions of the p-values from the hypothesis tests of non-monotonicity in trends. The null hypothesis is that increases in mortality among a lower-educated quartile has been no larger than increases of a higher-educated quartile within the same age/sex/race group. For each group, there are 6 hypothesis tests (Q1 vs. Q2, Q1 vs. Q3, Q1 vs. Q4, Q2 vs. Q3, Q2 vs. Q4, Q3 vs. Q4). Each CDF plots the results of 240 tests (40 groups x 6 tests per group). p-values are calculated via bootstrapping using 10,000 repeated samples within groups.

Figure 6. CDFs of p-values of tests of non-monotonicity in mortality trends comparing quartiles 1 and 2 to quartile 3



Note: Figure plots the cumulative distribution functions of the p-values from the hypothesis tests of non-monotonicity in trends between the bottom 3 quartiles. The null hypothesis is that increases in mortality among a lower-educated quartile has been no larger than increases of a higher-educated quartile within the same age/sex/race group. For each group, there are 2 hypothesis tests (Q1 vs. Q3, Q2 vs. Q3). Each CDF plots the results of 80 tests (40 groups x 2 tests per group). p-values are calculated via bootstrapping using 10,000 repeated samples within groups.

Appendix A Conceptual Framework

For simplicity, assume there are two groups: those with low socioeconomic status (SES) and those with high SES, proxied in our analysis by rank in the educational distribution. The stock of latent (unobserved) health capital at the end of a period for members of these two groups, denoted by H_l and H_h, is $\mu_l + \varepsilon$ and $\mu_h + \varepsilon$, for $\mu_h = \mu_l + \lambda$, with $\lambda > 0$ and ε a random variable normalized without loss of generalization as ε~(0,1). Death occurs if health capital falls below a threshold level H^0. For group j, where $j = \{l, h\}$, this occurs if:

$$H_j \leq H^0 \quad or \quad \mu_j + \varepsilon \leq H^0.$$

Defining $X = H^0 - \mu_l$, the probability of death for the low and high SES groups are:

$$R_l = F(X) \tag{A.1a}$$

and

$$R_h = F(X - \lambda) \tag{A.1b}$$

where $F(\bullet)$ is the cumulative distribution function of the relevant distribution. High SES individuals have lower death rates than their low SES counterparts: $R_h < R_l$, since $\lambda > 0$.

We begin by considering the effects of a health shock, S, that has a uniformly negative effect on the health capital of all individuals. The new health capital for members of group j is then $H_j = \mu_j - \mathrm{S} + \varepsilon$ and the risk of death of low and high SES persons becomes

$$R_l' = F(X + S) \tag{A.2a}$$

and

$$R_h' = F(X + S - \lambda). \tag{A.2b}$$

The probability of death has risen for both groups; however, what we are interested in are the absolute and relative changes in these risks. The absolute changes are:

$$\Delta R_l = R_l' - R_l = F(X + S) - F(X) \tag{A.3a}$$

and

$$\Delta R_h = R_h' - R_h = F(X + S - \lambda) - F(X - \lambda). \tag{A.3b}$$

The absolute change in the probability of death will then be higher for low SES group members if $\Delta R_l > \Delta R_h$, which occurs if:

$$F(X + S) - F(X + S - \lambda) > F(X) - F(X - \lambda). \tag{A.4}$$

This condition holds as long as the density of F is increasing on the interval to the left of H_0 and the shock is "small" (such that $S < \mu_l - H^0$). Intuitively, this occurs if deaths are "left-tail" events and the shock is not so large as to change this.

Next, consider the conditions under which the high SES group has higher *relative* increases in mortality than the low SES group, even though the absolute increase is smaller. This occurs by definition if $\frac{\Delta R_L}{R_L} < \frac{\Delta R_H}{R_H}$ or:

$$\frac{F(X+S)-F(X)}{F(X)} < \frac{F(X+S-\lambda)-F(X-\lambda)}{F(X-\lambda)} \tag{A.5}$$

Rearranging the inequality yields:

$$F(X-\lambda)F(X+S) < F(X)F(X+S-\lambda) \tag{A.6a}$$

or

$$\frac{F(X-\lambda)}{F(X)} < \frac{F(X+S-\lambda)}{F(X+S)} \tag{A.6b}$$

This expression holds for $S > 0$ if the density of F is increasing on the interval $[-\infty, H_0]$ which will occur for left-tail events with many common distributions (e.g. normal and T-distributions), although it need not be the case for others.

The preceding discussion suggests that equal size deleterious shocks to health capital will raise the absolute death rates more for low than high SES groups but that the opposite could be true for relative changes. However, there are additional reasons why equal shocks may also have larger relative effects as well for low SES individuals. For instance, assume the shocks are to the inputs into a Grossman-type health production function, where health capital at time t is determined by investments, occurring during the period and an exogenous depreciation rate (Grossman 1972). These investments are determined by a set of positive inputs that have diminishing marginal returns. A deleterious shock can be characterized as a reduction in the size of one of the (positive) inputs. Under this scenario, there are at least two reasons why equal size negative health shocks will have larger harmful effects on health investments and health capital for the disadvantaged. First, since high SES persons are likely to be further out on the "flat of the curve", where the marginal productivity of health investments is relatively low, a given negative investment shock will reduce health capital by less for them than for persons with lower initial levels of health capital, for whom the productivity of health investments is greater. Second, high SES individuals are likely to be more able to undertake compensatory investments to offset the negative effects of the shock. For example, when faced with increased pollution levels, they may have more ability to make local moves to less polluted areas or to make other investments to partially pollution-proof their residence. Similarly, they are likely to have greater access to high-quality medical care useful for mitigating the consequences of pollution.

The analysis of positive health shocks is largely the reverse of that just described, with greater absolute mortality reductions anticipated for the less advantaged and with unclear predictions for relative decreases in death rates. To the extent that improvements in medical technology or other similar positive health shocks are the norm, we expect to observe larger absolute (although again not necessarily relative) mortality improvements for disadvantaged groups with relatively elevated initial death rates.

Appendix B. Construction of Death Rates by Education Quartiles

This Appendix details our methods for constructing death rates by education quartile. Section B1 first describes the details of our imputation procedure to calculate deaths and population counts by years of education and demographic characteristics. These counts are then used to construct death rates by group, and represent the data used in estimating the main regressions of the paper as described in Section 3. Section B2 then describes the procedure to age-standardize death rates by education quartile, which follow common practices from the literature.

B1. Procedure to Estimate Population and Deaths by Years of Education

To estimate death counts (the numerator in the death rate calculations), we sum all deaths for the specified cell by age, gender, educational attainment, year and race/ethnicity, using *MCOD* data. We drop approximately 4,200 observations with missing age out of over 47.7 million recorded deaths during this period. Prior to 2003, information on single year of education is provided on the death certificates. Beginning in 2003, approximately 16 percent of deaths measure education in one of seven categories: 8th grade or less, 9-12th grades without a diploma, high school, some college (no degree), bachelor's degree, master's degree, or a doctorate/professional degree. By 2007, just over half of records classify education using these coarser groups, rather than single year of education, and in 2017, nearly all deaths are recorded using the seven education categories. For some classifications, we can reasonably assign a single year of education. Specifically, we treat high school graduation as 12 years of education, some college or associate's degree as 14 years, a bachelor's degree as 16 years, and a master's or doctorate/professional degree as 17 years of schooling. However, for the other education categories ("<= 8th grade" and "9-12th grade, no diploma"), this assignment cannot be done, since these broader categories include people with substantially different years of education; therefore, we develop an imputation procedure to use in these cases.

To implement the procedure, we first calculate the fraction of single year educational attainment, when these are provided, comprising each of the broader categories. For example, for deaths corresponding to grades 9 to 12 without a diploma, we calculate the percentages of deaths occurring among persons where the death certificate specifies 9, 10 and 11 years of education, respectively (and not just the broader education category). We then regress the percentages for each of these years of education on a quadratic trend in years and a full set of age, sex, and race/ethnicity interactions, with the sample restricted to those in the specified broader education categories (e.g. 9 to 12 years of education without a diploma). To ensure a large enough sample to make these extrapolations, we use wider than five-year age bins, specifically, combining those 25-39, 40-54 and 55-74 years of age. We restrict the time period for these regressions to be prior to and including 2010, since after that year fewer than 30 percent of deaths record single year of education. Next we use these estimates to predict the probability of persons with information only on the broad education category having the particular number of years of education, conditional on the three age aforementioned categories, age, sex, race/ethnicity and year of death.

A potential threat to this imputation strategy is that states adopting the broad education categories might have different distributions of within-category educational attainment to those that did not. To examine whether this was a problem, we first classified states according to whether they predominantly reported continuous years of education in 2010 versus those that primarily used the broader education categories. We then compared the distribution of deaths across these two classifications for those with 9, 10 and 11 years of education prior to 2003 (when all states used continuous education), conditional on having between 9 and 12 years of education (without a high school degree). We found that the distributions were nearly identical across the two types of states. We repeated this for those with 8th grade or less, and found similarly that the distribution of 0 through 8 years of education in the pre-2003 period was very similar between states that used different classification methods in 2010. These results suggest that the educational distributions in earlier years provide a useful indication of the predicted distributions in later ones.

Educational attainment is missing for roughly 5 percent of death certificates. We assume that the education distribution within a given year, race, sex, 5-year age bin is the same for these missing certificates as when education is reported, and include such deaths in the calculations using this allocation.

To estimate population counts, the denominator of death rates, we begin by assigning the number of years of education completed corresponding to the level of educational attainment, as measured in the *ACS*. While information on education is available from the 2000 Census, our analysis suggested that these data were not fully consistent with those reported in the ACS. Since we also use the ACS for other years, we choose to exclusively use the ACS to maintain comparability over time. This procedure is straightforward for categories up to grade 12 starting in 2008, since these are measured in single year bins. Prior to 2008, grades below 8^{th} grade were combined (nursery school to 4th grade, 5^{th} and 6^{th} grade, and 7^{th} and 8^{th} grade). We split these cases into each of the possible grades based on the distribution within a given race, sex and wide age bin among years 2008-2017. We record "no schooling completed", "nursery school, preschool", and kindergarten" as 0 years of education. We assume a high school degree is equivalent to 12 years, classify 12^{th} grade without a diploma as 11 years of schooling, and less than one year of college as 12 years. We assign “1 or more years of college credit, no degree” or an associate’s degree as 14 years of schooling and assume that a college degree without additional education is equivalent to 16 years. Education beyond a college degree is coded as 17 years of education. Using *ACS* sample weights, we then calculate the distribution of education measured from 0 to 17 years (excluding 13 or 15 years) by 5-year age categories, gender, survey year (and sometimes race/ethnicity). Finally, we multiply these population shares by the SEER population data corresponding to the age, gender, year and usually race/ethnicity cells to estimate population counts by single year of education and demographic sub-group.

It is important to acknowledge the assumptions implied by proportionately assigning deaths across quartiles using our methods. Novosad and Ravkin (2019) note that the proportional assignment, which is also used by Meara, Richards, and Cutler (2008) and Bound et al. (2015), treats mortality rates as being flat within education bins, and only allows for changes discretely across bins. By contrast, their methods assume a continuous latent education rank distribution, with mortality rates weakly declining in this rank. Assuming a step-function of mortality with proportional assignment is undesirable when education bins are wide, but the assumption is less problematic when education is measured in single years of schooling, as in our analysis. Novosad and Ravkin (2019) consider four education bins (less than high school, high school, some college, and bachelor's degree or higher), while we split education into 16 bins (0, 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 11, 12, 14, 16 and 17 years, where 14 includes all those with more than a year of college but who did not graduate and 17 includes all with at least one year of post-graduate education). Given the finer granularity of our measure of educational attainment, we view the assumption of constant mortality rates within single year of education as reasonable and potentially advantageous to analyses that divide the sample into just four education categories.

B2. Aggregation of Death Rates and Education Quartiles

The procedures described above result in education-quartile specific death rates calculated for demographic subgroups within 5-year age bins. In computing overall death rates for the aggregate group of 25-74 year olds (separately by sex), we adjust for changes over time in the age and education distributions over time by constructing weights for each 5-year age group a in education quartile i, based on 2017 population shares as:

$$w_{ai}^{2017} = \frac{pop_{ai}}{\sum_a pop_{ai}} \,, \tag{B1}$$

where pop_{ai} is the 2017 age group population for educational quartile i and $\sum_a pop_{ai}$ is the total population of that quartile. These aggregations are done based on age and education but not race groups, so the r subscript is not included. We then take the weighted average of death rates across age groups, standardized based on the 2017 age distribution:

$$\widehat{mort}_{it} = \sum_a (w_{ai}^{2017} \times mort_{ait}), \tag{B2}$$

where $mort_{ait}$ denotes the death rate for age-group a in education quartile i and year t, and $\widehat{mort}_{it}$ denotes the corresponding overall age-adjusted death rate for quartile i in that year.

To describe broad changes in the education distribution, we similarly also aggregate across groups. Specifically, we calculate quartile-specific average education as:

$$\widehat{educ}_{it} = \sum_a (w_{ai}^{2017} \times educ_{ait}), \tag{B3}$$

where $educ_{ait}$ denotes average years of education for age-group a in education quartile i and year t, and $\widehat{educ}_{it}$ indicates the corresponding overall age-adjusted average education for the quartile and year.

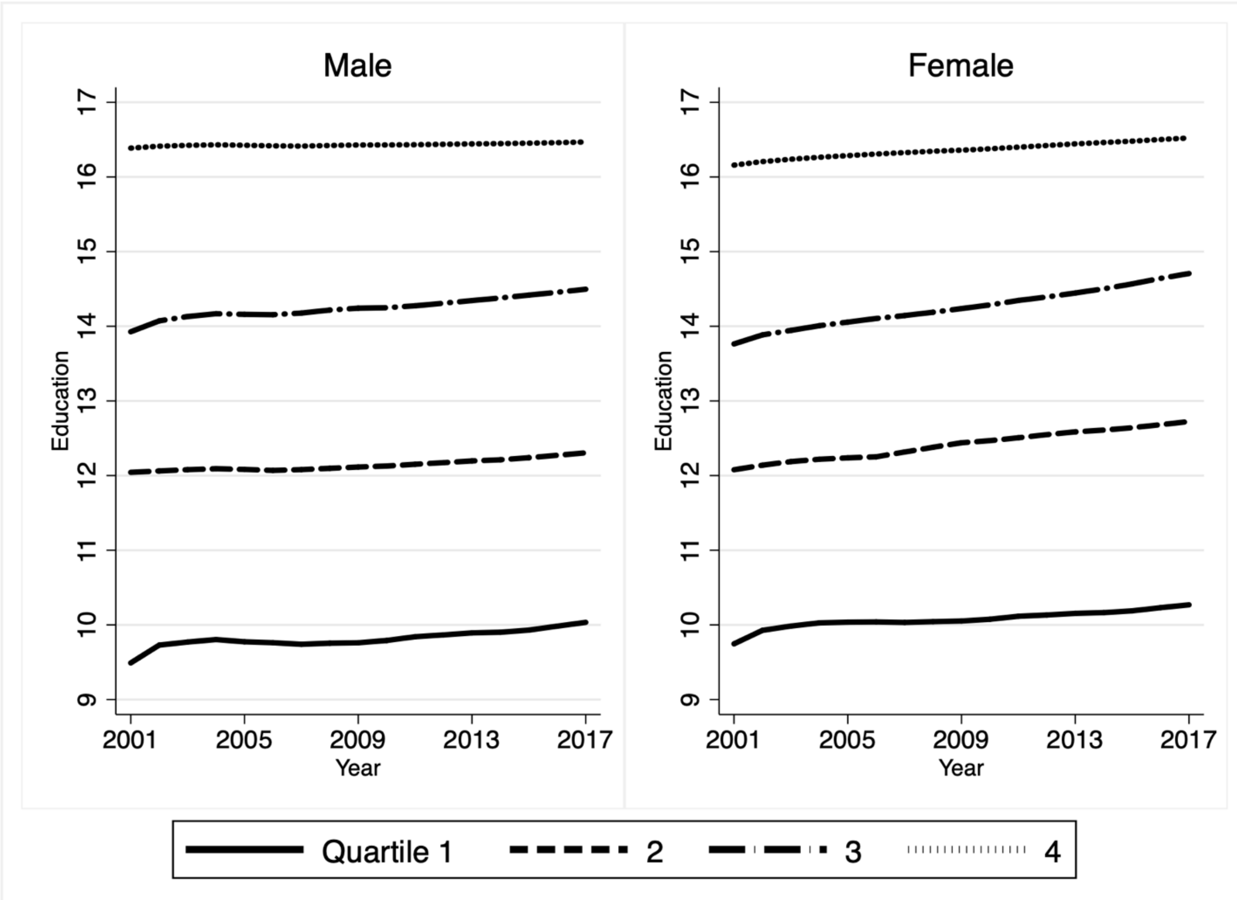
Appendix C: Additional Tables and Figures

Table C1. Number of violations of education-based monotonicity in mortality trends

Group	Males		Females	
	Log Death Rate	Death Rate	Log Death Rate	Death Rate
Any Violation (max = 40)	30	31	16	25
Type of Violation				
Q1 < Q2	10	7	4	4
Q1 < Q3	17	17	3	7
Q1 < Q4	0	9	1	10
Q2 < Q3	22	22	10	15
Q2 < Q4	2	16	3	17
Q3 < Q4	3	14	5	16

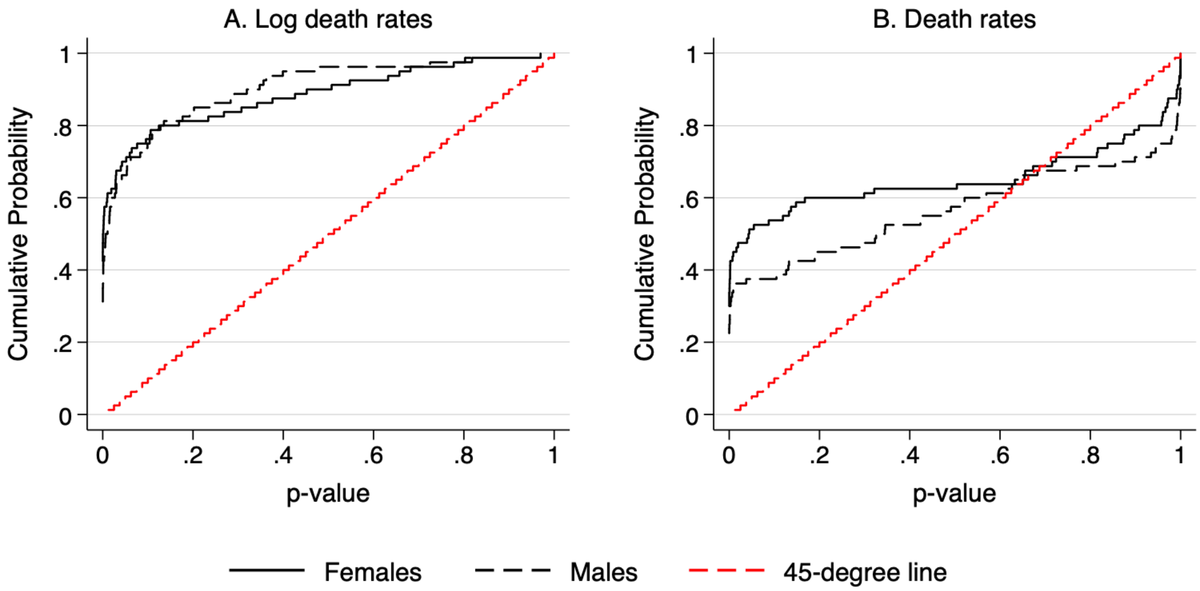
Note: Table shows the number of age-race-sex groups with monotonicity violation, defined to occur when a lower education quartile has slower mortality growth or larger decline than does a higher quartile for the same age, race and sex. This includes cases where Q1 has a better outcome than Q2, Q3 or Q4; Q2 has a better outcome than Q3 or Q4; or Q3 has a better outcome than Q4. The numbers in the row Q1 < Q2, for example, denote the number of cases when the mortality rate rose faster or fell more slowly for the first quartile than the second quartile of the same age-race-sex group.

Figure C1. Mean years of education by quartile and sex: 25-74 year olds



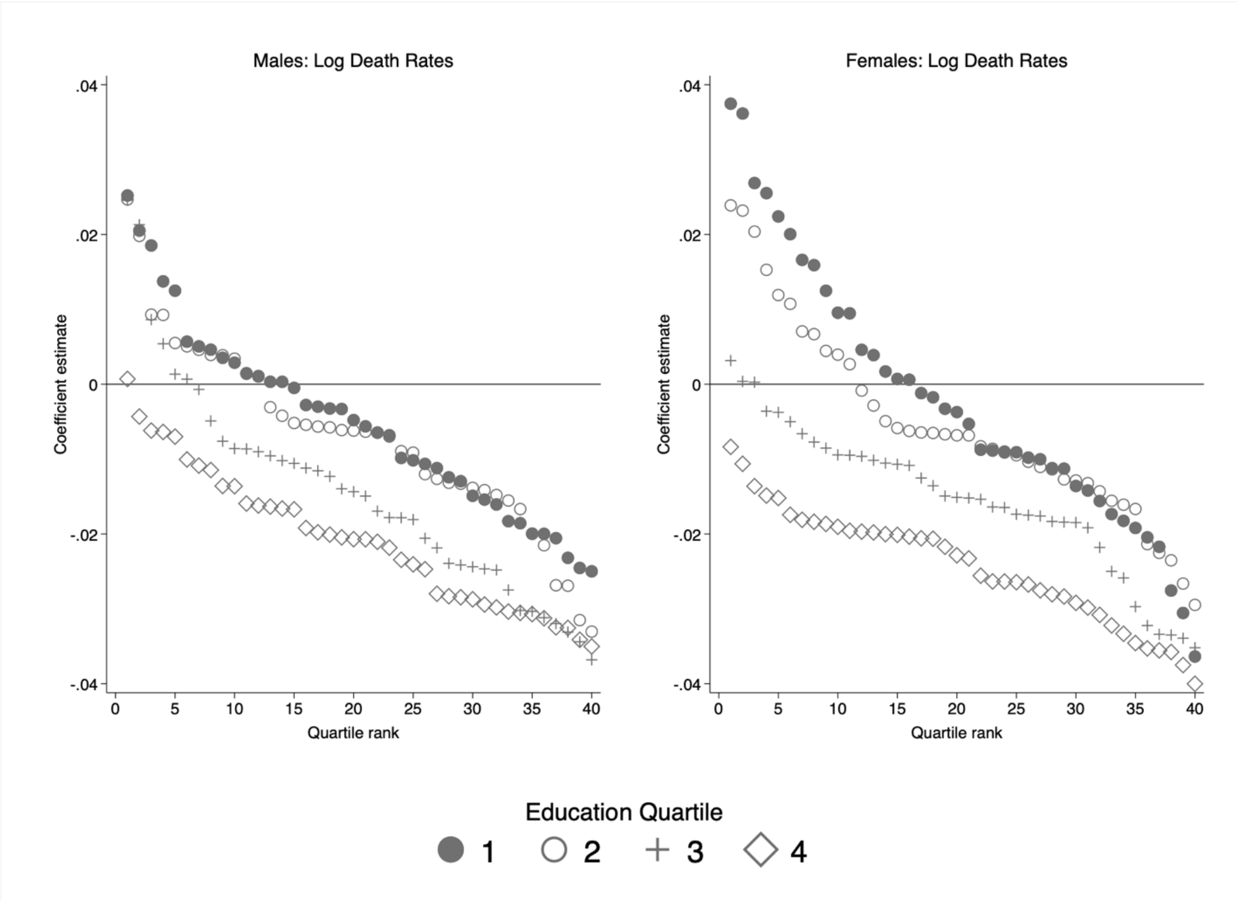
Note: Figure shows the average number of years of completed education by education quartile and sex for 25-74 year olds from 2001-2017. Education quartiles are calculated separately by sex, 5-year age group, and year using data from the ACS and SEER as described in Appendix B. The 2017 age distribution specific to each quartile and sex is used to age-standardize the mean years of education in the 25-74 year old group across time.

Figure C2. CDFs of p-values of tests of non-monotonicity in mortality trends comparing quartiles 1 and 2 to quartile 4



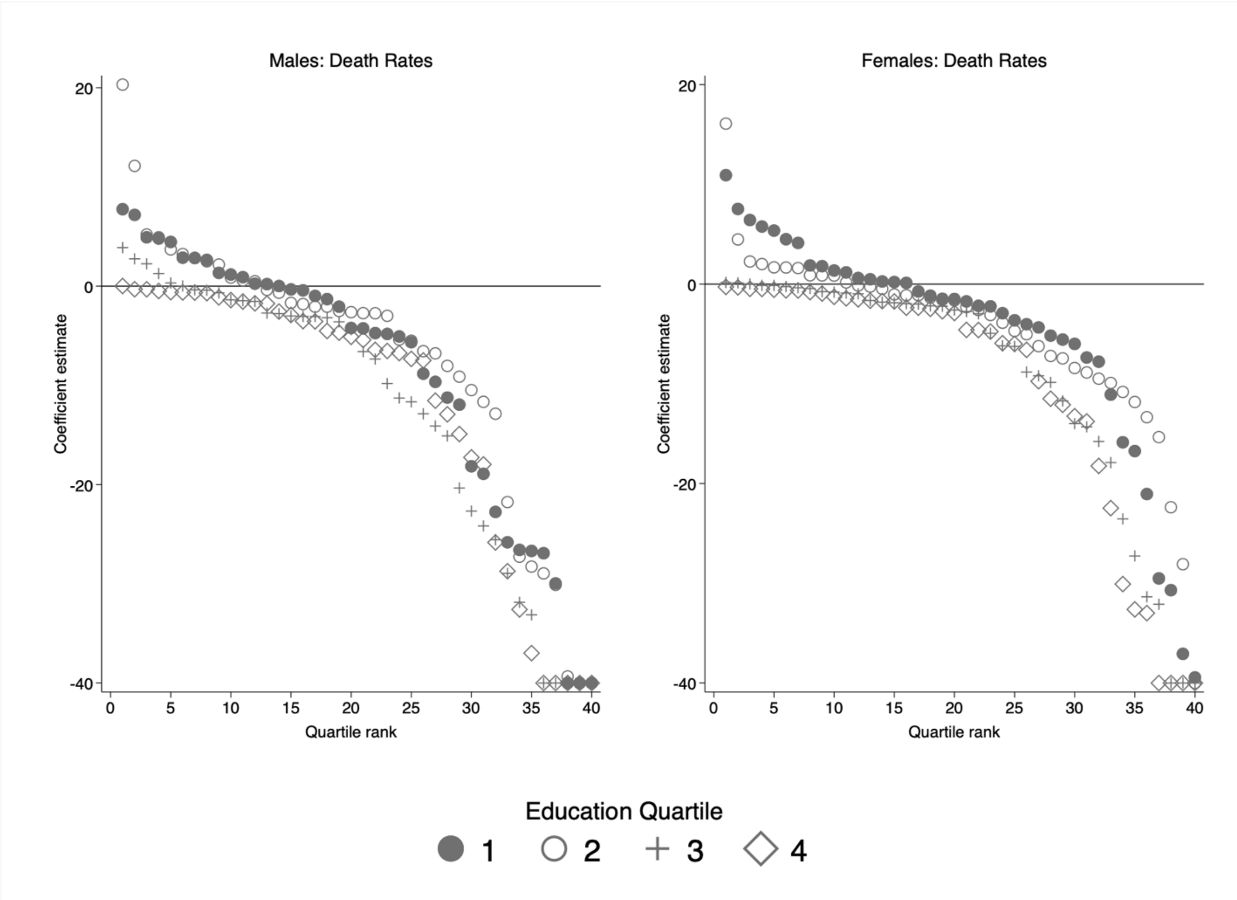
Note: Figure plots the cumulative distribution functions of the p-values from the hypothesis tests of non-monotonicity in trends between Q1 or Q2 and Q4. The null hypothesis is that increases in mortality among a lower-educated quartile has been no larger than increases of a higher-educated quartile within the same age/sex/race group. For each group, there are 2 hypothesis tests (Q1 vs. Q4, Q2 vs. Q4). Each CDF plots the results of 80 tests (40 groups x 2 tests per group). p-values are calculated via bootstrapping using 10,000 repeated samples within groups.

Figure C3. Ranked log death rate trends by race-specific education quartile



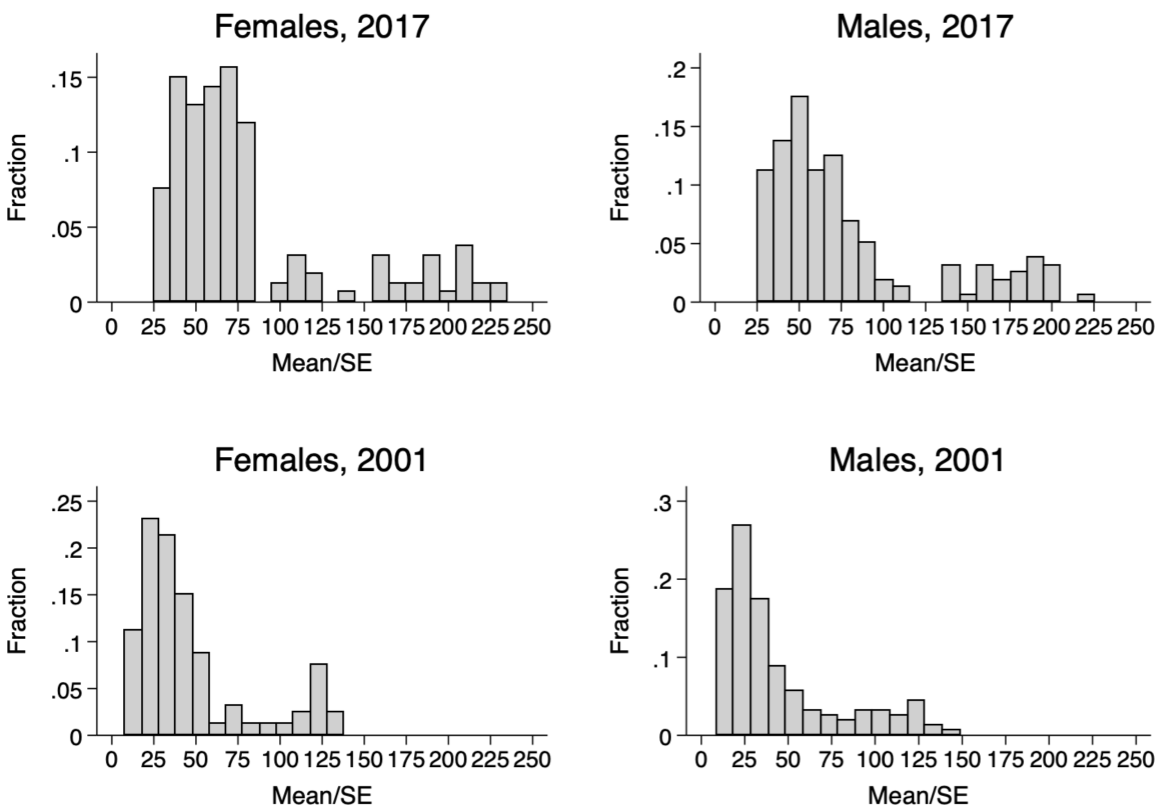
Note: Figure shows quantile plots of the trend coefficient estimates on log death rates from estimating equation (3), in which education quartiles are constructed separately for each race. Coefficient estimates are reported separately by education quartile, pooling races and ages together. The *x*-axis lists the rank of the trend coefficient estimate within each education quartile, in which higher ranks denote larger reductions in mortality rates.

Figure C4. Ranked death rate trends by race-specific education quartile



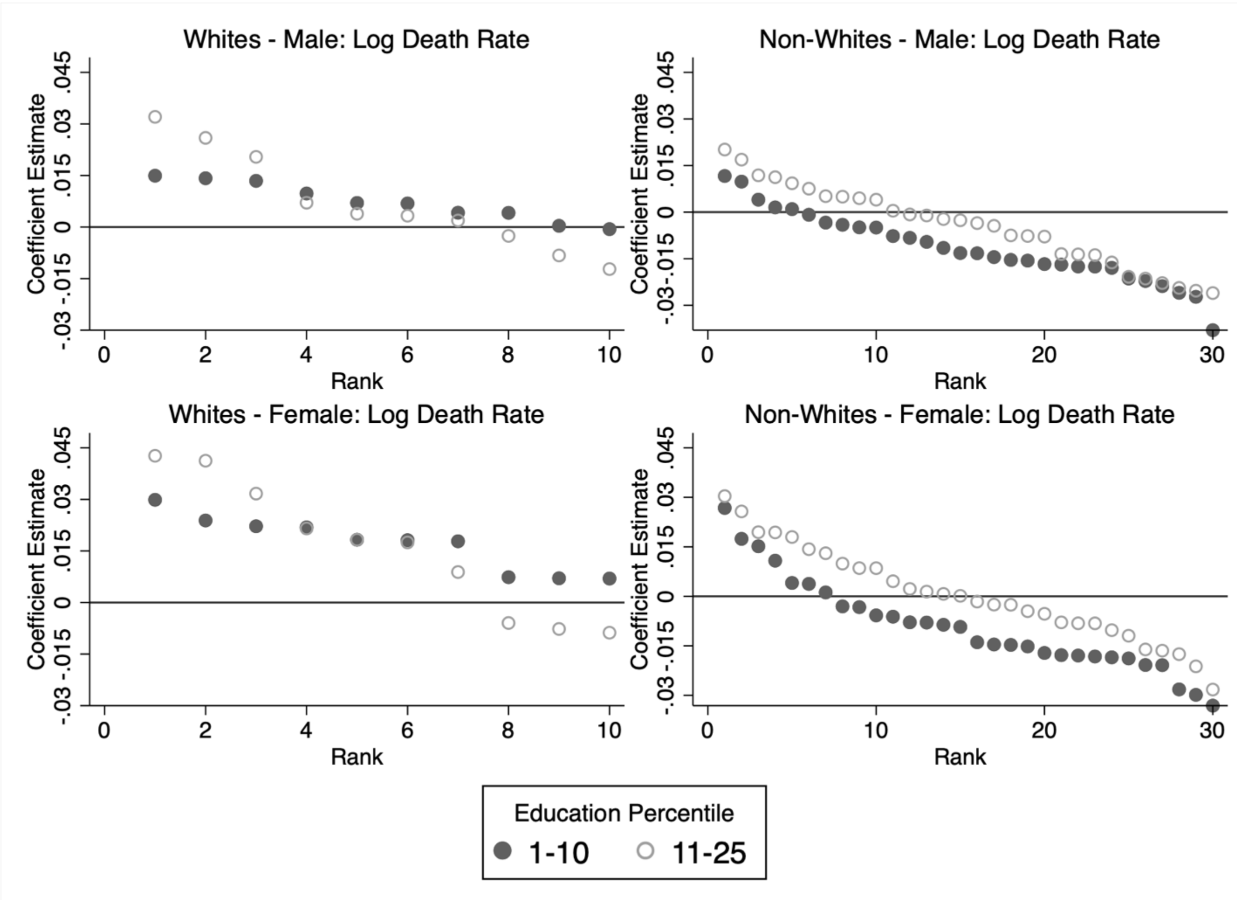
Note: Figure shows quantile plots of the trend coefficient estimates on death rates from estimating equation (3), in which education quartiles are constructed separately for each race. Coefficient estimates are reported separately by education quartile, pooling races and ages together. The x-axis lists the rank of the trend coefficient estimate within each education quartile, in which higher ranks denote larger reductions in mortality rates. For clarity, we bottom-code decreases in death rates at negative 40 to permit visible detection of differences for most of the distribution.

Figure C5. Histograms of Mean Education Shares Relative to Standard Error by Group



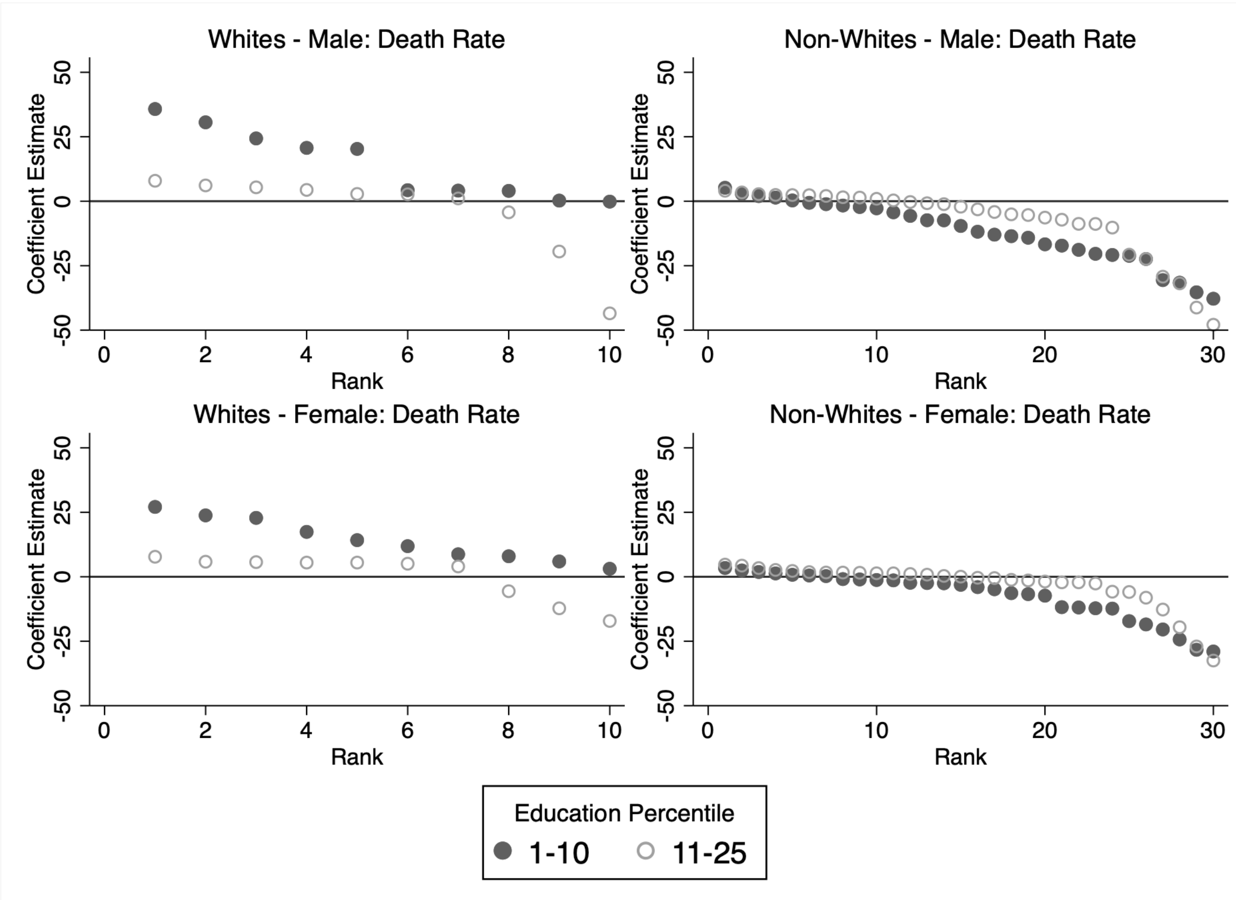
Note: Figure plots histograms of the coefficient of variation (Mean/SE) on the education shares from the ACS for each group of race, 5-year age, and four categories of education (less than high school, high school, some college, college grad or higher), separately by sex in 2001 and 2017. Bin width equals 10. Each histogram includes 160 groups. The smallest CV in 2001 is 8 for Hispanic women aged 70-74 with a college degree (mean = 6.3%, SE = 0.7%). The largest CV in 2017 is 234, for white women aged 30-34 with a college degree (mean = 47.8%, SE = 0.2%).

Figure C6. Ranked log death rate trends among lowest quartile, by race



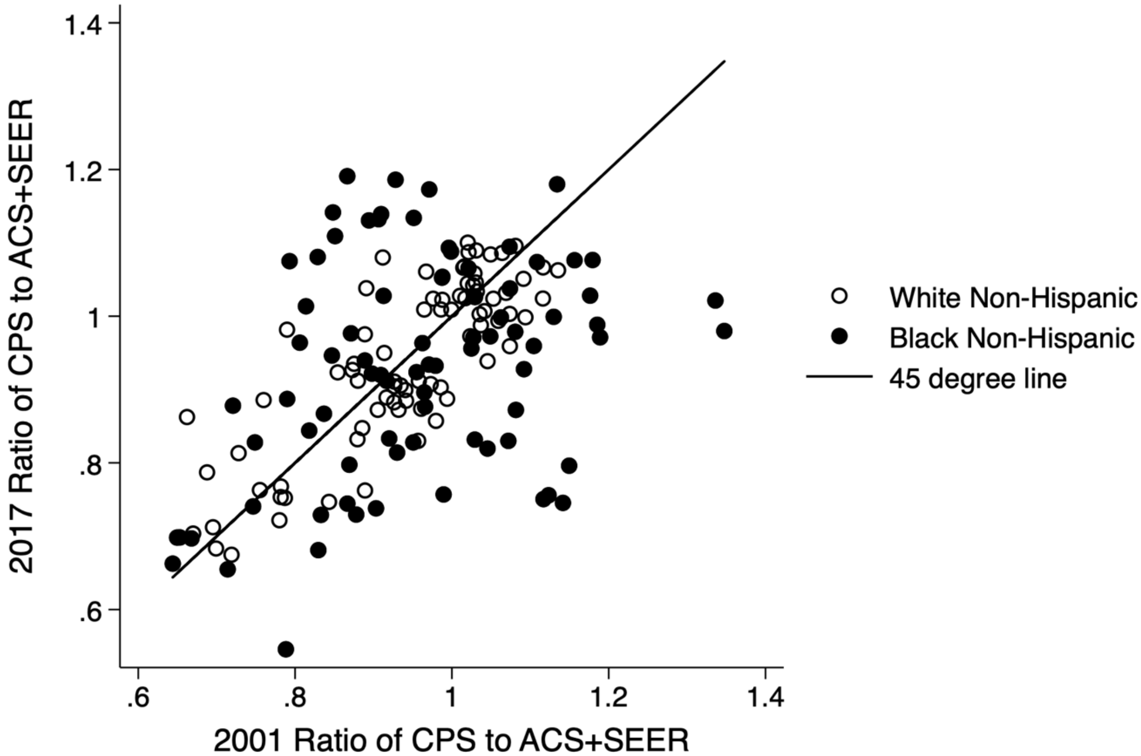
Note: Graph plots distributions of the trend coefficient estimates on log death rates that split the bottom quartile into the lowest 10 percent (hollow circles) and the 11^{th} to 25th percent (shaded circles) of the education distribution. The y-axis lists the trend coefficient estimates on log death rates. The x-axis lists the rank of the trend coefficient estimate within either the bottom 10 percent or 11^{th}-25^{th} percentiles, in which higher ranks denote larger reductions in mortality rates.

Figure C7. Ranked death rate trends among lowest quartile, by race



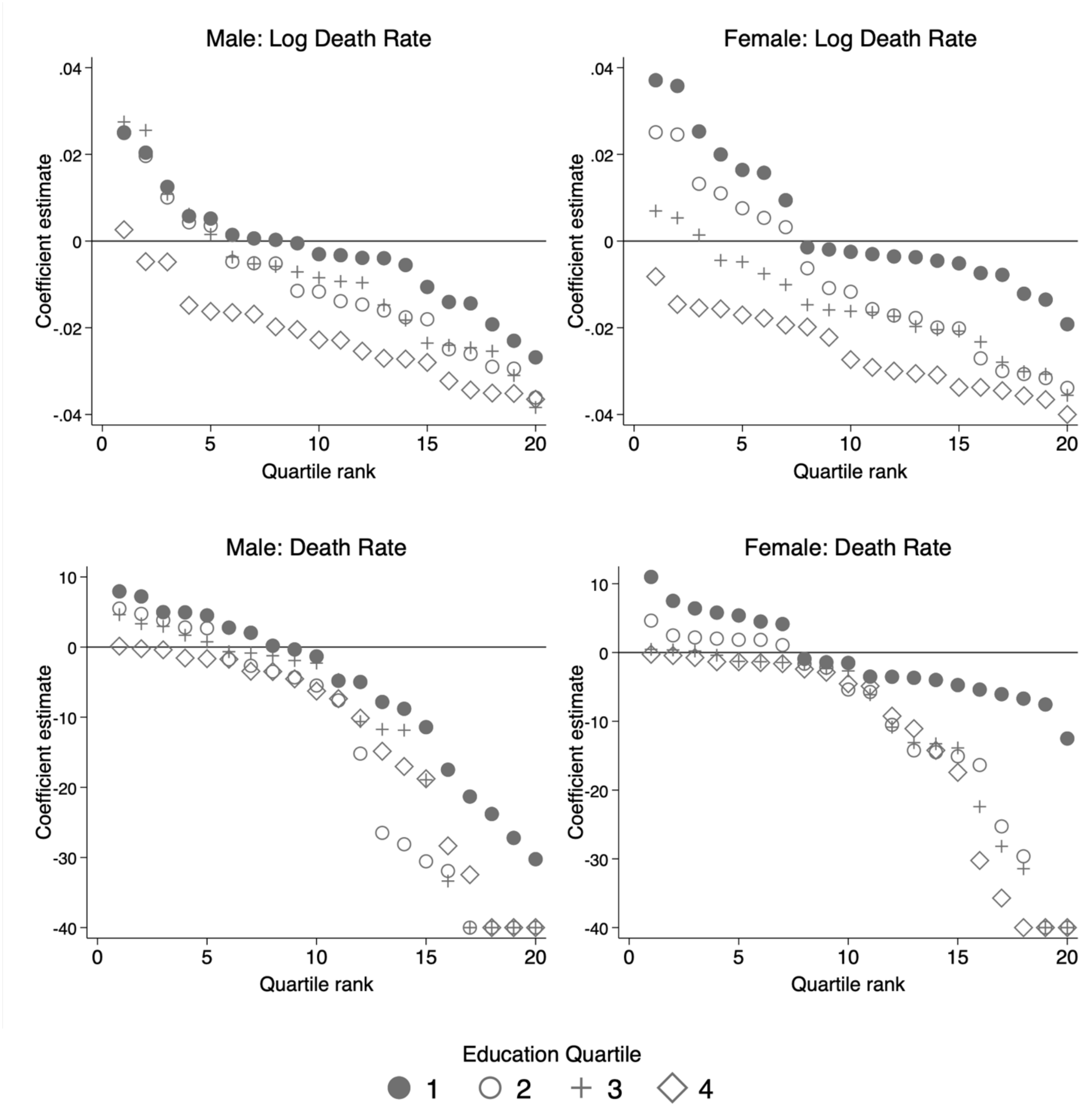
Note: Graph plots distributions of the trend coefficient estimates on log death rates that split the bottom quartile into the lowest 10 percent (hollow circles) and the 11th to 25th percent (shaded circles) of the education distribution. The y-axis lists the trend coefficient estimates on death rates. The x-axis lists the rank of the trend coefficient estimate within either the bottom 10 percent or 11th-25th percentiles, in which higher ranks denote larger reductions in mortality rates.

Figure C8. Comparison of Population Estimates using CPS vs. ACS and SEER



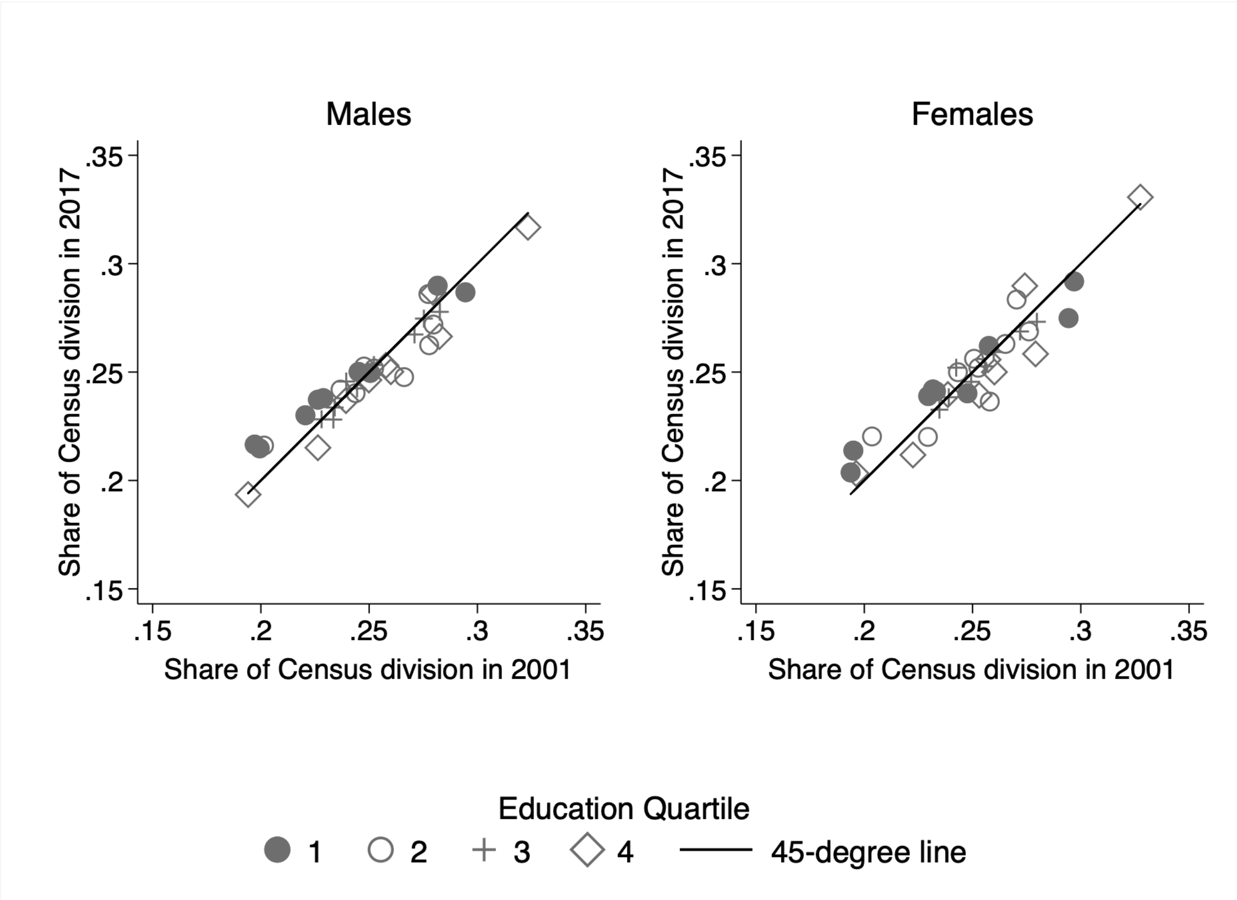
Note: Figure plots the ratio of population estimates from the CPS to those estimated from the combination of the ACS and SEER for 5-year age bands, sex, race, and four education categories (Less than high school, high school, some college, college). The *y*-axis displays the ratio using 2017 data and the *x*-axis displays the ratio using 2001 data. A ratio of 1.2 is interpreted as the population estimated from the CPS is 20 percent larger than that estimates by multiplying the SEER by the share of that demographic cell in the ACS.

Figure C9. Ranked mortality trends by education quartile, whites and blacks only



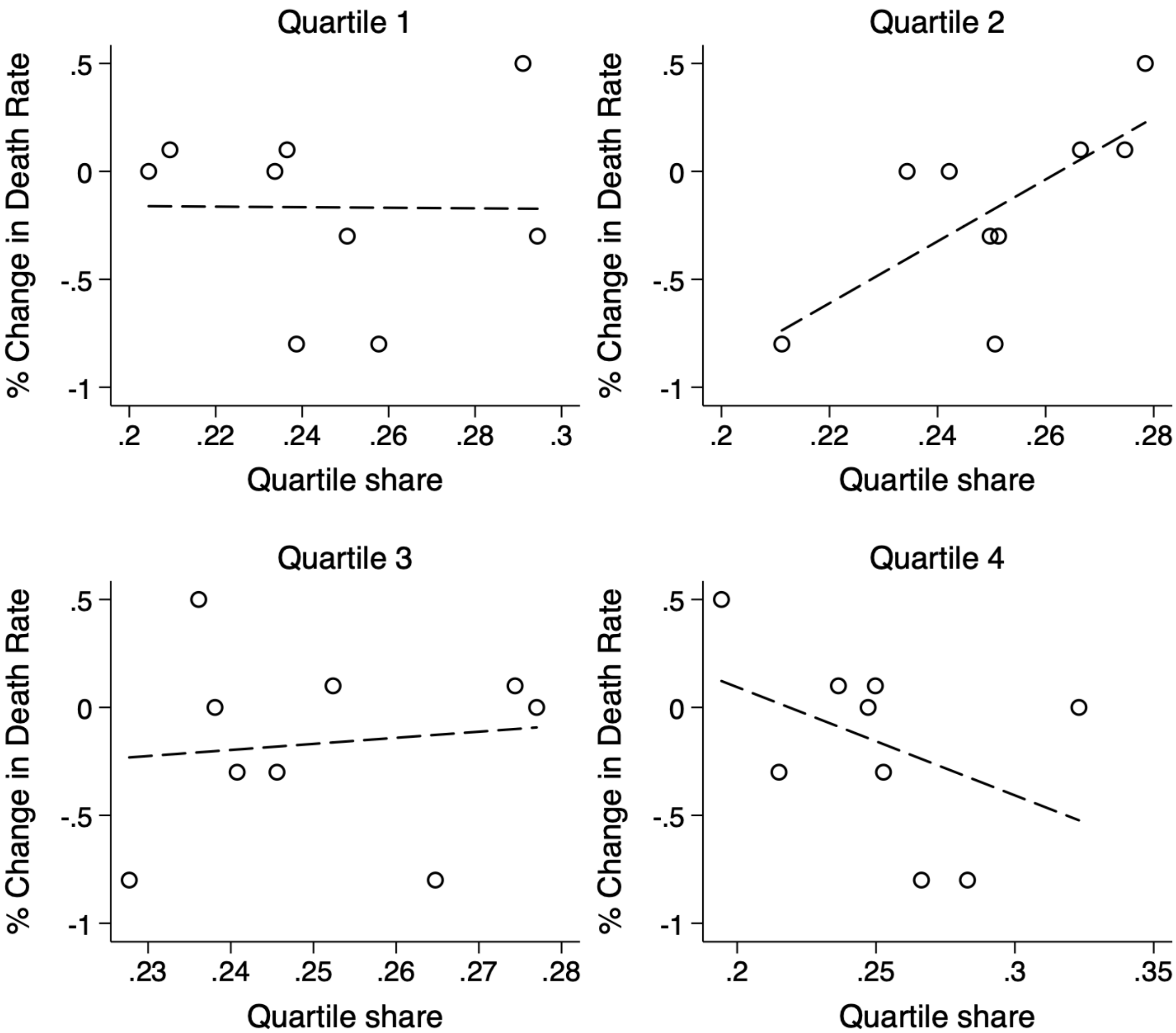
Note: Figure shows quantile plots of the trend coefficient estimates on log and levels of death rates from estimating equation (3), with only whites and blacks in construction of quartiles. Coefficient estimates are reported separately by education quartile, pooling races and ages together. The *x*-axis lists the rank of the trend coefficient estimate within each education quartile, in which higher ranks denote larger reductions in mortality rates. For clarity, we bottom-code decreases in death rates at negative 40 to permit visible detection of differences for most of the distribution.

Figure C10. Shares of Education Quartiles within Census Divisions, 2001-2017



Note: Figure shows the fraction of education quartiles within Census divisions in 2001 and 2017, calculated separately for males and females. Each observation corresponds to a quartile within a particular Census division. The solid line denotes the 45-degree line. Points located farther from the 45-degree line indicate larger charges in the fraction of the Census division composed of that particular quartile. Points located close to the 45-degree line, indicate little change over time in the composition of quartiles within regions. We use the national thresholds for quartiles calculated using the ACS and the SEER as described in the text and used throughout the analysis.

Figure C11. Percentage Change in Death Rates vs. Quartile Shares within Census Divisions



Note: Figures plot the relationship between the average annual percentage change in census region death rates against the average share of each education quartile. The mortality data are from Figures 9, 10, and 11 in Woolf and Schoomaker (2019), with males and females combined. The average education quartile shares within Census regions are calculated from 2001 to 2017, and also combine males and females because the shares vary little within Census divisions. We use the national thresholds for quartiles calculated using the ACS and the SEER as described in the text and used throughout the analysis. The relative change in the death rates, within Census regions, is negatively related to the share of population in Q4 positively related to the share in Q2. There is little relationship between mortality changes and the share within Q1 or Q3.

Appendix D: Education Quartiles vs. Categories

Our main analysis divides education groups into quartiles, so as to examine a constant proportion of the educational distribution over time. The more commonly used method is instead to use fixed educational categories. When doing so, secular increases in education will cause the proportions of the population in the categories to change, with increases in higher and decreases in less educated categories. This introduces potentially serious selection biases and a potential tradeoff between simplicity and accuracy. It is an empirical question whether the benefits of the more complicated strategy we follow are worth the additional complexity. For this reason, this appendix examines whether the results obtained substantially differ when using education quartiles rather than categories.

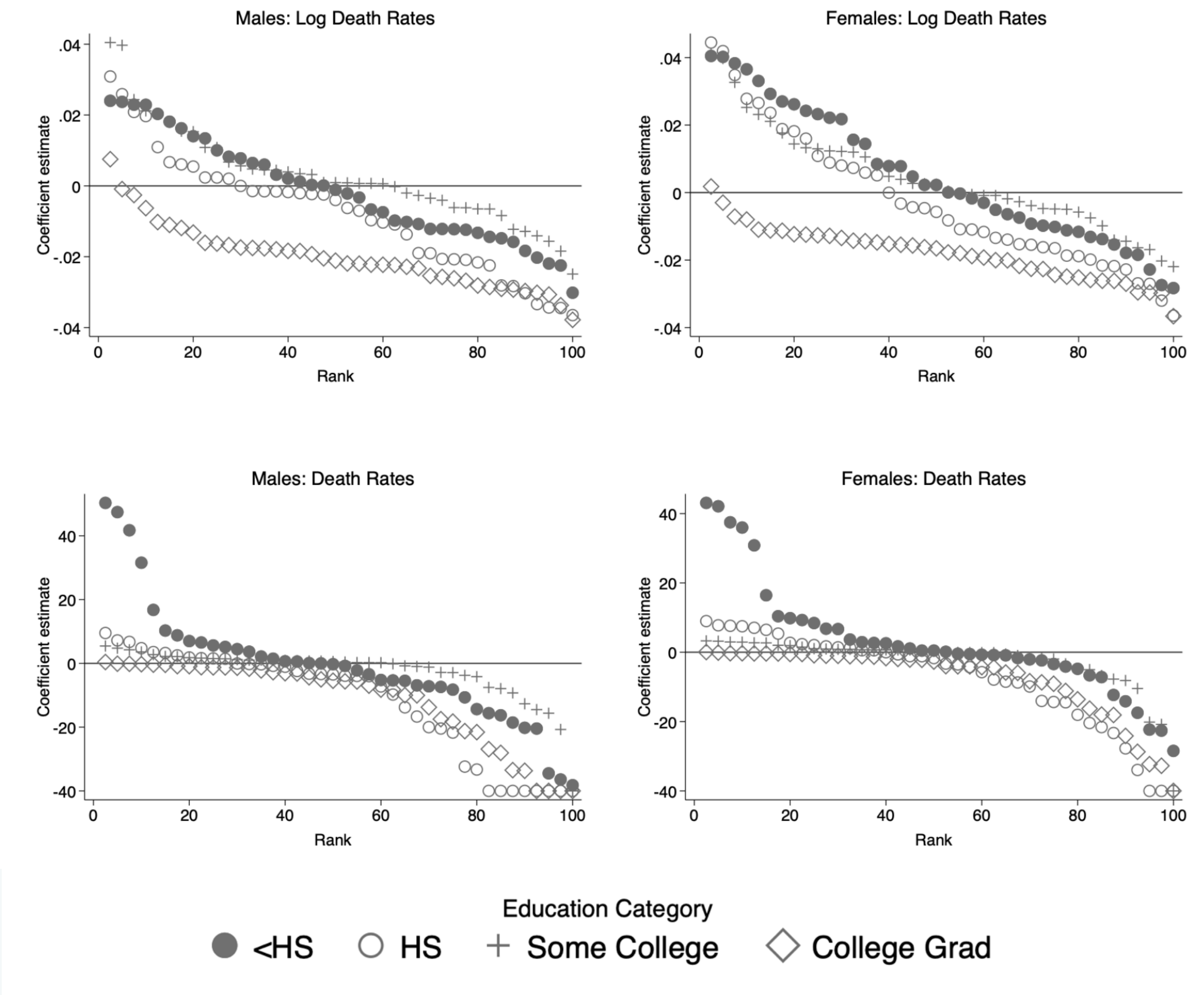
The four education quartiles can be roughly matched to the following four schooling categories less than high school graduate; high school graduate but no college, some college but no degree, and college degree or more. These categories approximately correspond to the overall average educational attainment of Q1-Q4, although quartile-specific years of schooling vary across groups and generally increase over time.

A positive correlation between categorical and quartile-based mortality trends is obtained, as would be expected, although with considerable variation across groups. For example, the correlations are 0.90 and 0.71 when examining male logs and levels of death rates, and 0.89 and 0.72 for corresponding female outcomes. The rank correlations for males range from 0.83 to 0.88, compared to 0.84 to 0.88 for females. The categorical and quartile trend coefficients have different signs in around 16 percent of cases for men and for 19 percent of groups for women, although this mostly occurs when the absolute value of the trend coefficient is small.

To provide a better indication of the importance of the sensitivity of the results to the use of educational categories versus quartiles, Appendix Tables D1 and D2 expand on some of the prior analysis. In each case, the original estimates for quartiles are shown first, followed by the corresponding results using education categories. Using education categories often results in misclassification of both the worst and best performing groups. Between 2 and 5 of the 10 age-race-education quartiles with the largest trend increases in death or log death rates are misidentified when using corresponding education categories, with particularly poor performance (half of the 10 groups misidentified) for levels of death rate (see Table D1). In addition, the use of categories leads to a substantial overstatement of the magnitude of the rise for the worst-off groups. For instance, the death rates of white male Q1 aged 55-59 and 30-34 were estimated to increase by 9.5 and 7.0 per 100,000 annually, the two largest increases of any of the 160 groups. By comparison, the estimated increases for same aged white males with less than high school education were 31.6 and 8.8 per 100,000 annually.

However, these estimates are erroneous, reflecting increasing negative selection into these groups as educational attainment rose over time. Similarly, the largest increases for females were the 12.0 per 100,000 annual rise in death rates estimated for 50-54 year old Q1 whites. However, once again the growth for corresponding aged white women with less than a high school education was over two and a half times as large: 30.8 per 100,000. The estimated increase for white women aged 65-69 with less than a high school degree was 43.1 per 100,000, but the rate a decline of 6.7 per 100,000 for Q1 white women of this age. Table D2 shows that the use of education categories, rather than quartiles, also frequently misidentifies the best performing groups, although both the magnitudes and identification of groups in the top 10 are much closer to those obtained using quartiles.

Figure D.1 Ranked log and level death rate trends by education category



Note: Figure shows quantile plots of the trend coefficient estimates on log and levels of death rates from estimating equation (3), using education categories rather than quartiles. Coefficient estimates are reported separately by education category, pooling races and ages together. The *x*-axis lists the rank of the trend coefficient estimate that is specific to each education quartile, ranging from largest increase to largest reduction.

Table D1. Groups with largest mortality increases by education quartiles and categories

	Log Death Rate								Death Rate							
	Quartiles				Categories				Quartiles				Categories			
Rank	Race	Age	Educ	Coef	Race	Age	Educ	Coef	Race	Age	Educ	Coef	Race	Age	Educ	Coef
Males																
1	W	30	3	0.029	W	25	3	0.040	W	55	1	9.46	W	65	1	50.34
2	W	30	2	0.025	W	30	3	0.040	W	30	1	7.01	W	60	1	47.45
3	W	30	1	0.024	W	30	2	0.031	W	50	1	6.57	W	70	1	41.76
4	W	25	3	0.022	W	25	2	0.026	W	30	2	5.87	W	55	1	31.58
5	W	25	2	0.021	W	35	3	0.024	W	55	3	5.52	W	50	1	16.77
6	W	25	1	0.018	W	30	1	0.024	W	55	2	4.71	B	65	1	10.29
7	O	25	1	0.016	O	30	1	0.024	W	35	1	4.65	W	30	2	9.53
8	O	30	1	0.016	O	25	1	0.023	W	25	1	4.43	W	30	1	8.80
9	W	35	2	0.013	W	25	1	0.023	W	25	2	4.18	W	35	2	7.21
10	W	35	1	0.012	O	30	3	0.021	O	50	2	3.92	W	25	1	7.00
Females																
1	W	30	1	0.036	W	30	2	0.044	W	50	1	11.95	W	65	1	43.10
2	W	25	1	0.034	W	25	2	0.042	W	55	1	8.29	W	70	1	42.11
3	O	25	1	0.028	W	30	3	0.041	W	45	1	7.11	W	55	1	37.49
4	W	30	2	0.028	W	25	1	0.040	W	30	1	5.81	W	60	1	35.98
5	W	35	1	0.025	O	25	1	0.040	W	35	1	5.61	W	50	1	30.84
6	W	25	2	0.025	W	25	3	0.040	W	50	2	5.31	W	45	1	16.45
7	O	30	1	0.024	W	30	1	0.038	W	40	1	4.96	W	40	1	10.40
8	W	50	1	0.021	W	50	1	0.037	W	25	1	4.04	W	35	1	9.83
9	W	45	1	0.018	W	35	2	0.035	W	45	2	2.98	W	30	1	9.30
10	W	35	2	0.017	W	55	1	0.033	W	30	2	2.93	W	50	2	9.00

Note: This table shows the 10 groups with the largest estimated increases in logs or levels of death rates, separately by sex and whether education quartiles or categories are used in estimating equation (3). Education categories also range from one to four and refer respectively to less than high school graduate, high school graduate without college, some college but no degree, and a Bachelor's degree or more.

Table D2. Groups with largest mortality decreases by education quartiles and categories

	Log Death Rate								Death Rate							
	Quartiles				Categories				Quartiles				Categories			
Rank	Race	Age	Educ	Coef	Race	Age	Educ	Coef	Race	Age	Educ	Coef	Race	Age	Educ	Coef
Males																
160	B	70	3	-0.039	B	50	4	-0.038	B	70	3	-161.1	B	70	2	-174.2
159	B	50	4	-0.036	H	45	2	-0.036	B	70	4	-101.8	H	70	2	-105.9
158	B	70	4	-0.035	B	45	2	-0.034	B	70	2	-90.8	B	65	2	-91.0
157	B	45	4	-0.035	B	50	2	-0.034	B	65	3	-87.8	B	70	4	-90.4
156	H	65	4	-0.035	B	45	4	-0.034	B	65	2	-87.8	O	70	2	-58.3
155	B	45	2	-0.034	H	40	2	-0.033	H	70	3	-82.3	H	65	2	-53.8
154	H	70	3	-0.034	B	70	4	-0.031	W	70	3	-69.5	W	70	4	-53.7
153	B	40	4	-0.033	H	70	2	-0.030	W	70	4	-63.4	B	50	2	-50.2
152	H	45	2	-0.033	B	40	4	-0.030	H	70	4	-53.5	H	70	4	-47.1
151	H	40	2	-0.033	O	65	1	-0.030	B	65	4	-50.5	W	70	2	-46.5
Females																
160	B	45	4	-0.040	B	45	4	-0.037	B	70	3	-115.5	B	70	2	-123.4
159	B	60	4	-0.036	H	70	2	-0.036	B	70	4	-69.9	H	70	2	-75.0
158	B	70	3	-0.036	B	70	2	-0.032	B	65	3	-62.8	B	65	2	-68.6
157	H	70	3	-0.036	B	25	4	-0.030	B	70	2	-61.9	B	70	3	-43.9
156	W	65	4	-0.035	B	60	4	-0.030	H	70	3	-61.3	B	70	4	-42.2
155	B	25	4	-0.033	B	50	4	-0.030	B	65	2	-58.6	H	65	2	-34.0
154	B	65	4	-0.033	O	65	1	-0.028	B	65	4	-49.6	W	70	4	-32.7
153	W	70	4	-0.033	O	60	1	-0.027	W	70	4	-48.7	B	65	4	-32.3
152	O	65	3	-0.032	B	65	2	-0.027	O	70	3	-39.6	B	60	4	-28.7
151	O	70	4	-0.031	B	55	4	-0.027	B	60	4	-35.8	O	70	1	-28.4

Note: This table shows the 10 groups with the largest estimated decreases in log or levels of death rates, separately by sex and whether education quartiles or categories are used in estimating equation (3). Education categories also range from one to four and refer respectively to less than high school graduate, high school graduate without college, some college but no degree, and a Bachelor's degree or more.