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Has mortality risen disproportionately for the least educated? ☆

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ABSTRACT

We examine whether the least educated population groups experienced the worst mortality trends at the beginning of 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 mortality between the top 25 percent and the bottom 75 percent is growing. However, there are many groups for whom these patterns are reversed – with better experiences for the less educated – or where the differences are statistically indistinguishable.

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 even before the famous Whitehall studies (Marmot et al., 1984; Marmot et al., 1991), with many analyses conducted in recent years.[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 mortality. Contemporary studies have also highlighted differences in mortality trends over time by educational group (Meara et al., 2008; Cutler et al., 2011; Montez et al., 2011; Olshansky et al., 2012; Montez et al., 2019a). 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 (Murphy et al., 2018; National Center for Health Statistics, 2017), 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 to 2013.

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[1] Seminal research in the 19th century includes work by Louis R. Villermé and William Farr (Buck et al., 1995; Whitehead, 2000; Julia and Valleron, 2011), and in the 20th century by Kitagawa and Hauser (1973). Bosworth (2018) provides an overview of difficulties in conducting this research, and some recent results.

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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, analyzing mortality at fixed percentiles of the distribution of educational attainment will 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 (hereafter simply "race"), and five-year age ranges. We then use linear regression to quantify how mortality trends vary from 2001 to 2017 across education quartiles. For clarity of exposition, we hereafter use the terms "death rates" or "log death rates" to distinguish between absolute versus relative changes and "mortality" to indicate more general experiences that are consistent across both measures.

We find that the relationship between mortality trends and education varies sharply by gender. For women, the magnitudes of the mean reductions in relative death rates have generally increased monotonically with education. Conversely, males in the three lowest education quartiles often have fairly similar experiences. The gap in mortality between the top 25 percent and the bottom 75 percent is growing for both men and women. However, these averages conceal substantial heterogeneity across groups. We document that these patterns are frequently reversed within sex-age-race groups, such that lower education quartiles experienced larger mortality reductions than their more highly educated counterparts. These findings are robust to a number of extensions and sensitivity checks.

Ours is not the first analysis to attempt to adjust for changes in the distribution of education over time. Bound et al. (2015) employ somewhat 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 improvements. Goldring et al. (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 et al. (2020) develop a partial identification approach to bound changes in mortality rates between 1992 and 2018. They argue that mortality increases are concentrated in the bottom decile of the educational distribution for both blacks and whites. Conversely, we find that adverse mortality experiences extend throughout the bottom quartile for women and the lower three quartiles for men, on average, but with considerable heterogeneity across groups.

Our results provide a more complete picture of recent mortality trends among the full U.S. population than the studies just cited by including Hispanics and other races in addition to whites and non-Hispanic blacks (hereafter "blacks"), and using recent data, which is important given rapidly changing mortality trends.[4] We also supply a comprehensive analysis of the frequency with which mortality trends are not monotonic in education, and we demonstrate that the results obtained using fixed education categories (e.g., high school graduates), rather than education quartiles, will often be incorrect in important ways. We believe that our approach to addressing compositional change provides results that are complementary to other recent work on this topic, corroborating certain patterns and providing new insights on how the gradient between education and mortality has changed differently for demographic groups. Better understanding these trends and heterogeneity of the patterns across groups is of direct interest for those concerned about disparities in health, and also has implications for the incidence of programs like Medicare and Social Security that are linked to longevity.

2. Construction of death rates

We calculate educational quartiles separately by 5-year age bins, gender, race and year. For compactness, we frequently abbreviate the first through fourth education quartiles as Q_1 through Q_4, where Q_1 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 r, sex s and education quartile i, in year t as:

$$M_{arist} = \frac{D_{arist}}{P_{arist}} \tag{1}$$

where D_{arist} and P_{arist} refer, respectively, to the number of deaths and population of the subgroup. We compute these rates for five-year age groups ranging from 25 to 74. The youngest group analyzed are aged 25–29 because completed education is not meaningfully measured below this age. Our oldest group are aged 70–74 because mortality selection becomes an increasing issue at older ages:

[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 4, we also consider quartiles that are both race- and sex-specific.

[4] Other studies examining aspects of these same issues have also excluded some population groups and cover an earlier time period. For example, Hendi (2015) examines how shifting education distributions have affected education gradients in life expectancy using data for non-Hispanic whites only, from 1991-2005. Sasson (2016) uses data from 1990-2010 and excludes Hispanics and other nonwhites in his analysis of education category differences in both mean values of life expectancy and variation around these averages.

those who survive into their 80s and 90s, for example, are less frail than those who die earlier, which confounds comparisons between cohorts at different ages (Vaupel et al., 1979).

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.[5] Information on total population by age, gender, race, and year is obtained from the National Cancer Institute's *Surveillance Epidemiology and End Results* (*SEER*) database.[6] These are combined with information on the distribution of educational attainment by age, gender, race, and year using data from the Census Bureau's *American Community Survey* (*ACS*). We use 2001 as the starting period for our analysis because this is the first ACS year.

An empirical challenge in constructing death rates by education quartile is that schooling is measured in discrete units and the information provided is not fully comparable across time periods or data sources. With continuous education measures, this 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.[7] Obtaining 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 graduate). We address this limitation using an imputation procedure, detailed in Appendix A, to estimate death and population counts by single year of education.

As an overview, some states in the *MCOD* switch from reporting deaths by single year of education to using broader categories partway through the sample, while others continue to use single year of education for much of the sample period. We use the distribution of education among the latter set of states to predict single year of education in the former, using a regression that includes a quadratic trend and all interactions between age, sex, and race. We validate this strategy by comparing the distributions of single year of education when it is observed in both sets of states during early years of the sample. We find the distributions are nearly identical, which suggests that predictions from states continuing to measure education in single years provide a useful indication of the education distribution in states that switch to using categories earlier. Estimating population counts also requires an adjustment since the *ACS* switches from using categories to single year of education in 2008. Fortunately, the categories are fine enough that we can assign single year of education in most cases under reasonable assumptions, such as a high school degree is equivalent to 12 years of schooling and 12th grade without a diploma equals 11 years of schooling. Prior to 2008, schooling below 8th grade is separated into three categories (nursery school to 4th grade, 5th and 6th grade, and 7th and 8th grade), and we use the distribution of single year of education by age, sex, and race measured over the 2008–2017 *ACS* to separate these categories into single years of schooling. Appendix A provides additional description about these procedures.

The second complication is that a single year of education may straddle quartiles. For example, if 22 percent of the group have 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. In the example just provided, three-sevenths of deaths and population for the 11-year group are assigned to Q_1 and four-sevenths to Q_2.[8] Our empirical approach treats mortality rates as being constant within education bins, while changing discretely across bins. We view this assumption as reasonable given that we are measuring single years of education, which are quite granular.

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. Our primary analyses use general thresholds, reasoning that it does not make sense for two individuals of the same age, sex, and education to be placed in different quartiles just because their races differ.[9] One implication of general thresholds is that the lower quartiles will be disproportionately populated by race 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 4.

3. Mortality trends by education quartile

This section describes overall trends in educational attainment and death rates by educational quartile for ages 25–74 combined. We first construct aggregated measures of schooling and death rates in each year using common age standardization techniques described in Appendix A. Next, we examine how log death rate changes differ with education using data from the 160 age-race-education groups for each sex based on regression analysis.

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

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

[7] In two cases we combine groups. Persons with at least 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.

[8] Three-sevenths are assigned to Q_1 since $\frac{25-22}{29-22} = \frac{3}{7}$ and four-sevenths to Q_2 since $\frac{29-25}{29-22} = \frac{4}{7}$.

[9] 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. See also Kaestner et al. (2020) for how education and health may differ over the life-cycle and across birth cohorts.

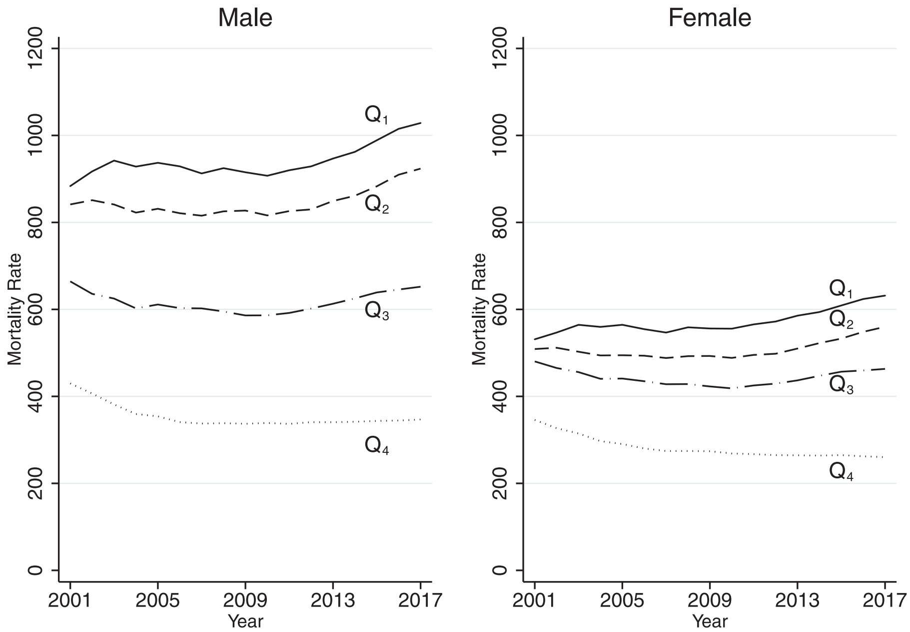

Fig. 1. Death rates by education quartile and sex, 25–74 year olds. Note: Figure shows age-adjusted death rates per 100,000 by education quartile among 25–74 olds from 2001 to 2017. Death rates are calculated using MCOD, ACS, and SEER data as described in Appendix B. Death rates are age-standardized using the 2017 age distribution specific to each education quartile. Education quartiles are calculated separately by sex, age and year. Q1 denotes the least educated quartile and Q4 denotes the highest educated quartile.

3.1. Descriptive trends

The educational levels of 25–74 year olds rose over time for all quartiles, with larger growth for females. The average increases were smallest for the highest-educated quartile, Q_4 (0.07 years for men and 0.35 years for women over the period), and greatest for the second-highest quartile, Q_3 (0.52 years for men and 0.90 years for women). These overall patterns conceal much larger increases for some groups. For example, educational attainment rose by 1.3 and 1.0 years for 60–64 year old Q_1 and Q_3 men, and by 1.2, 1.6 and 0.8 years for Q_1, Q_3, and Q_4 women of the same age.

The overall sex-specific and education quartile-specific annual death rates are shown for each year from 2001 to 2017 in Fig. 1. The gradient between educational attainment and average death rates is clear. Moreover, Q_4 is the only quartile to experience steady declines over time for both sexes. Death rates also fell for Q_3 males and females until 2010 but then rose. Trends for the bottom two quartiles were flat or slightly increasing until 2010, and then exhibited a sharper increase.

Fig. 2 stratifies the patterns by age, splitting into four groups for each sex: 25–39, 40–54, 55–64, and 65–74 years old. For both men and women, the greatest reductions in death rates occur for those over age 65, with large decreases observed among the two highest quartiles of the education distribution, Q_3 and Q_4. By contrast, the increase in death rates occurring in recent years is particularly apparent among the least educated quartiles for those aged 25–39 for both men and women. Rates have been flatter for 40–64 year olds, compared to those older or younger. Across ages and sexes, the highest education quartile has most consistently experienced declines in mortality.

3.2. Regression estimates

We now turn to regression methods to more carefully analyze the time series patterns summarized in Figs. 1 and 2. One challenge in describing mortality trends across education quartiles, age, and sex is that the relationship is not always linear. We pursue three complementary approaches to verify that our qualitative results are not driven by specific modeling decisions. First, we estimate linear trends for education quartiles over two time periods—2001 to 2009 and 2009 to 2017. We choose these periods because they are of equal length (split at the midpoint of our analysis period) and because the evidence above suggests that mortality trends are roughly linear during each sub-period. This strategy is similar to fitting a spline with a knot in 2009. Second, we fit linear trends to the full sample period, which provides a single trend estimate for each age-sex-race education quartile. This provides a concise summary and uses all the available data, but with the potential cost of imposing a functional form that may be insufficiently flexible to accurately describe the patterns for all groups. Third, we measure mortality changes from the beginning of the sample period to

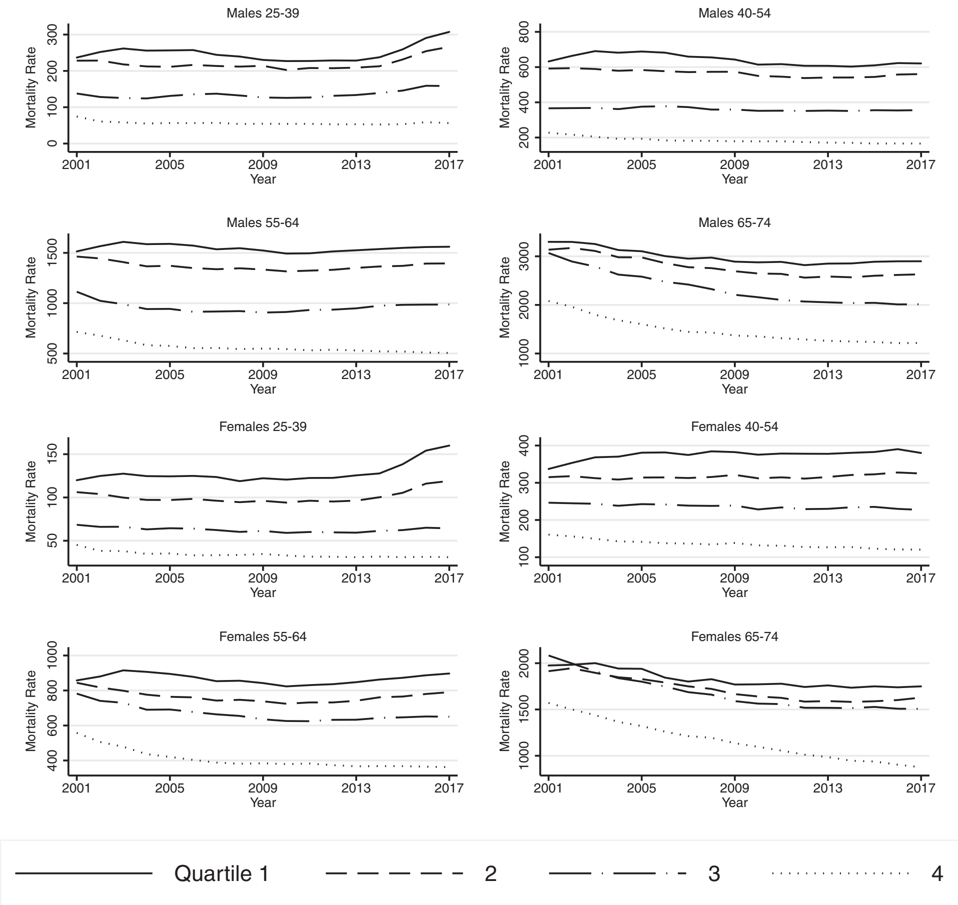

Fig. 2. Death rates by education quartile, age, and sex. Note: Figure shows age-adjusted death rates per 100,000 by education quartile among by age group and sex from 2001 to 2017. Death rates are calculated using MCOD, ACS, and SEER data as described in Appendix B. Education quartiles are calculated separately by sex, age and year.

the end for each group. In calculating these changes, we average rates from the first and last three years (2001 to 2003 and 2015 to 2017) to reduce noise from a single year. This approach has the advantage of not imposing any functional form assumptions, but at the cost of not using information from the majority of years. Collectively, the three strategies yield a comprehensive description of how mortality has changed from 2001 to 2017.

While the descriptive mortality trends in Fig. 1 standardize for changes in the age distribution of the population, changes in death rates are likely to be dominated by groups with high baseline death rates, such as older individuals. For this reason, our initial regression estimates focus on log death rates, under the rationale that percentage changes are less subject to this issue.

Specifically, we analyze how log mortality death rates have changed differentially over time across education quartiles by estimating:

$$M_{arist} = \beta_{aris} + \pi_s \times t + \pi_{is} \times t + \epsilon_{arist} \quad (2)$$

where β_{aris} is a group-specific fixed-effect; $\pi_s \times t$ is the linear time trend for the reference group, Q_1; and $\pi_{is} \times t$ where i can take the values 2, 3, and 4, is the trend for Q_2, Q_3, and Q_4 relative to the lowest quartile of education Q_1. We estimate the regression models separately for men and women, so that each includes 160 age-race-education quartile groups. Standard errors are clustered by age, race and education, and we weight each cell by its population to obtain nationally-representative estimates.

We first estimate Eq. (2) separately for years 2001 to 2009 and for 2009 to 2017. The regression estimate of π_s, that is $\hat{\pi}_s$, is the trend "main effect" from the model. The point estimates for the trend of remaining groups are calculated as this main effect plus the quartile-specific trend coefficients. For instance, the estimated trend for Q_2 is $\hat{\pi}_s + \hat{\pi}_{2s}$. The 95 percent confidence intervals (CIs) are centered on the Q_2 through Q_4 total effects and indicate whether the corresponding trend is statistically significantly different from that for Q_1. This can be visually observed by whether the CI crosses the dotted horizontal line showing the estimated Q_1 trend.

Fig. 3 presents these estimates for each time period. Patterns across quartiles are different for men and women, but are qualitatively similar in both early and later years. Specifically, for men, the magnitudes are comparable for the bottom three education quartiles, Q_1 to Q_3, whereas there is evidence of more monotonic trends across education quartiles for women. The highest-educated quartile,

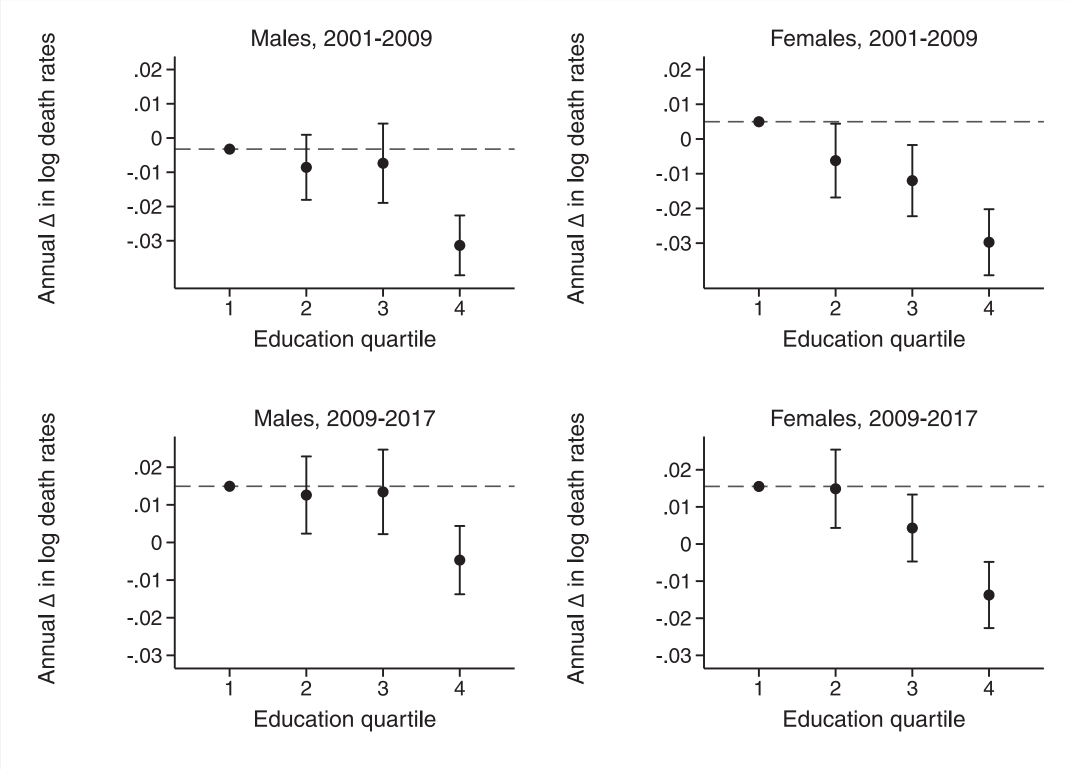

Fig. 3. Estimated Trend Differences by Quartile and Time Period. Note: Figure shows regression results from estimating Eq. (2) by time periods. Dots represent estimates on the trend coefficient for each quartile, with those for quartiles 2, 3 and 4 calculated by adding the estimate on the corresponding 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.

Q4, experienced better trends than other quartiles in both time periods and for both sexes. As suggested by Fig. 1, the magnitudes differ substantially between the time periods. Between 2001 and 2009, trends are negative for all but Q_1 females. Between 2009 and 2017, however, only the most-educated quartile of females experience declines. Similarly, Q_1 through Q_3 males experience little change in log death rates from 2001 to 2009, and increases in them from 2009 to 2017. By contrast, mortality improved for the highest quartile in the earlier period and with little change in the later one.

Fig. 4 presents corresponding results from estimating Eq. (2) over the entire 2001–2017 period. For men, the three lowest quartiles of education, Q_1 to Q_3, had virtually no change in mortality, with annual reductions of 0.1 percent or less in log death rates, while the most educated quartile, Q_4, had substantial reductions, averaging almost 1.5 percent per year. Conversely, for women, the higher the educational quartile, the better the evolution of mortality. We estimate that female death rates in Q_1 and Q_2 rose by 0.6 and 0.2 percent per year, respectively, compared to annual decreases of 0.6 percent and 1.9 percent for Q_3 and Q_4, respectively. Of particular importance for our analysis, the relative changes across education quartiles are almost identical when we compare the whole period, 2001–2017 with either of the two 9-year subperiods.

Finally, as our third approach, we estimate the following specification to measure the total change in log death rates over the entire sample period:

$$\Delta \ln\left(M_{aris}\right) \equiv \ln\left(M_{aris\tau}\right) - \ln\left(M_{aris0}\right) = \gamma_{ars} + \psi_{is} + e_{aris} \tag{3}$$

where $M_{aris\tau}$ is the mortality rate averaged between 2015 and 2017 for age group *a*, race *r*, quartile *i*, and sex *s*, and M_{aris0} is the corresponding average for 2001 to 2003. This specification includes a single observation for each age/sex/race/quartile. The ψ_{is} coefficients measure the mean sex-specific change for education quartile *i*. As above, we cluster standard errors by age and race, and weight each cell by its population.

The results of estimating Eq. (3), which are presented in Appendix Fig. B2, are remarkably consistent with Fig. 3, and again show strong evidence of monotonicity in changes by education for women. For men, the estimates for Q_2 and Q_3 are similar in magnitude, as they also were above, but are now lower than for Q_1 and statistically distinguishable. This result is driven by the high uptick in mortality at the end of the sample period for the least educated, previously identified in Figs. 1 and 2. Once again, the highest-educated quartile has mortality declines that are substantially larger than for the others.

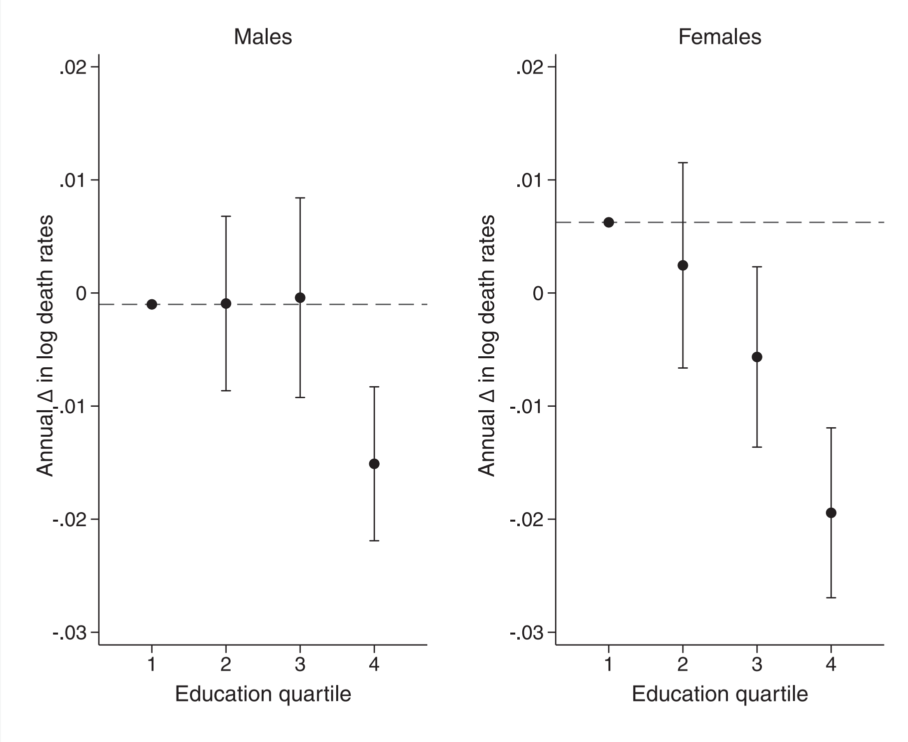

Fig. 4. Estimated Trend Differences by Quartile, 2001–2017. Note: Figure shows regression results from estimating Eq. (2). Dots represent estimates on the trend coefficient for each quartile, with those for quartiles 2, 3 and 4 calculated by adding the estimate on the corresponding regression trend interaction 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.

Taken together, we interpret these results to indicate that mortality patterns have differed sharply by gender. Men in the bottom three education quartiles have experienced roughly similar patterns, while for women, the gradient is widening across the education distribution. One common result across gender is that the gap in mortality between the top 25 percent and the bottom 75 percent is growing. We also demonstrate that while the specific estimates of mortality experiences vary over the three estimation strategies examined here, the relative performance of different education quartiles is measured fairly consistently across each approach. This finding is relevant for the analysis of heterogeneity below where we estimate additional models that use linear time trends covering the full sample period.

3.3. Heterogeneity

Up to this point, we have focused on average education quartile-specific gradients in mortality changes. However, these averages could conceal considerable heterogeneity across age and race groups. It is possible, for example, that mortality trends are worse for lower education quartiles than higher education ones for some groups but not for others. To identify how educational gradients have changed across groups, we estimate the following regression:

$$M_{arist} = \beta_{aris} + \delta_{aris} \times t + u_{arit} \tag{4}$$

where all variables are as previously described. Unlike Eq. (2), this specification includes separate time trends for all 320 age-race-/ethnicity education quartile groups (160 for each sex), without a trend main effect, and so provide information on changes in death rates for each group individually. The coefficients of primary interest, δ_{aris}, show the group-specific annual mortality changes. Of particular interest here is examining to what extent less-educated quartiles have experienced identifiably worse trends than higher-educated quartiles within the same demographic group.

We define death rate or log death rate trends to be *non-monotonic* if an education quartile has slower growth or a larger decline in mortality than a higher quartile for the same age, race, and sex. Thus a non-monotonicity occurs if: Q_1 does better than Q_2, Q_3 or Q_4; Q_2 does better than Q_3 or Q_4; or Q_3 does better than Q_4.

A simple comparison of the trend coefficient estimates reveals that non-monotonicities are common, occurring for 30 and 31 of 40 male groups for logs and levels of death rates, and for 16 and 25 of female groups (Table 1). For log death rates, the non-monotonicities

Table 1
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				
$Q_1 < Q_2$	10	7	4	4
$Q_1 < Q_3$	17	17	3	7
$Q_1 < Q_4$	0	9	1	10
$Q_2 < Q_3$	22	22	10	15
$Q_2 < Q_4$	2	16	3	17
$Q_3 < Q_4$	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 Q_1 has a better outcome than Q_3, Q_3 or Q_4; Q_2 has a better outcome than Q_3 or Q_4; or Q_3 has a better outcome than Q_4. The numbers in the row $Q_1 < Q_2$, 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, based on the linear trends from estimating Eq. (4) in the main text.

most frequently occur because Q_3 does worse than Q_1 or Q_2, while both Q_3 and Q_4 are often observed to have worse trends than Q_2 for death rates. Non-monotonic mortality patterns are much more common for nonwhites than whites. For males, 50% and 30% of groups of whites experience non-monotonic trends in levels and logs of death rates, compared to 83% and 93% of nonwhite groups. Non-monotonic trends are also common for nonwhite women, occurring for 53% and 83% of groups, but white females never exhibit non-monotonic changes in education gradients (see Appendix Table B2 for details).

To verify these results are not driven by the assumption of linear trends, we again calculate the difference from the beginning and end of our sample period using 3-year averages for each. Appendix Table B1 replicates the same comparisons as Table 1 and shows nearly identical counts in violations of monotonicity. This similarity reinforces the qualitative result of heterogeneity in mortality patterns by education level within quartiles.

These results just described demonstrate that for most groups, there is at least one instance where a lower education quartile has more favorable mortality trends than a higher one. However, even when they do not, the experiences of lower education quartiles may be only marginally worse. To formally examine these comparisons, we conducted 1-sided tests of the null hypothesis that a lower quartile has *equal or better* mortality trends than a higher education quartile; the alternative hypothesis is that the lower quartile has *worse* trends. Specifically, we use the group-specific trend estimates from Eq. (4) to compare: Q_1 vs. Q_2, Q_3 and Q_4; Q_2 vs. Q_3 and Q_4; and Q_3 vs. Q_4. We frequently fail to reject the null hypothesis that less-educated quartiles have equal or better than higher-educated quartiles.[10] This analysis confirms the frequency of relatively good mortality performance of lower quartiles compared to their more educated counterparts. Appendix B provides details of these tests and results.

4. Robustness and extensions

4.1. Education quartiles vs. categories

Our strategy of examining education quartiles, whose thresholds vary across groups and over time, is considerably more complicated than using fixed education categories (e.g., less than high school graduate, high school graduate, some college, college graduate). Here we summarize whether such extra effort is warranted by examining whether education quartiles produce meaningfully different results compared to fixed categories. Appendix C provides detail on this analysis and the findings.

The correlations between the quartile and categorical log death rate trends 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 and the evidence of monotonically declining educational gradients is much weaker for women. In fact, log death rates for both men and women with some college have risen by more than for those never attending college, though the differences are not statistically significant (Appendix Fig. C1). These comparisons by education category are confounded by compositional changes, unlike our preferred 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

[10] For example, using a *p-value* less than 0.10 as the criterion for rejecting the null hypothesis, we fail to reject for between 30% and 60% of comparisons depending on the sex and outcome (levels vs. logs of mortality rates) examined.

whites near retirement age experienced the largest growth in death rates with either classification method, but the use of educational categories overstates their increases by a factor of two or three (Appendix Table C1). 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.

4.2. Race-specific education thresholds

Our main specifications used age- and sex- but not race-specific thresholds for categorizing education quartiles. Appendix Fig. B6 reproduces the estimates from Fig. 4, allowing the quartile thresholds to also differ by race. There continues to be a monotonic pattern between education quartiles and mortality trends for females but it is slightly weaker than before. In contrast with Fig. 4, the point estimates of Fig. B6 show a clear educational gradient of male mortality trends, although differences for the bottom three quartiles are not statistically significant. We continue to find the greatest average declines among the highest-educated quartiles for both men and women.

4.3. 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 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.[11] Appendix Fig. B7 displays histograms of the mean divided by the standard deviation (i.e. the inverse of the coefficient of variation) of these shares by sex-age-race group. Most are estimated with a high degree of precision: the median inverse coefficient of variation is 33 in 2001 and 64 in 2017. The 5th percentiles are 13 in 2001 and 31 in 2017. The large inverse coefficients of variation provide support to our main results, and indicate that sampling variation in education shares is unlikely to change their interpretation.

4.4. Decomposing mortality experiences of the bottom quartile

We compare our results to recent research by Novosad et al. (2020) who use a bounding procedure to examine how mortality rates have changed for quantiles of the education distribution since 1992. They find the worst mortality experiences are concentrated among the lowest deciles of whites and blacks. To provide corresponding findings, we estimate models where the bottom quartile is split into those at or below the 10th percentile versus above it. In contrast to Novosad et al. (2020), we show similar trends for the bottom decile and the 11th to 25th quantile for both whites and blacks (Fig. B8). When including Hispanics and other races too, we find the 11th to 25th quantiles do slightly *worse* than the bottom decile. The disparities between our results and those of Novosad et al. (2020) might be attributable to differences in the data sources used to estimate the education distribution, the time periods analyzed, or the bounding procedure they employ. Other findings are more similar between the studies: both document that mortality increases extend to earlier ages and that using quantiles instead of categories can make important differences in some cases. Appendix B provides additional discussion that compares our methods and results in greater detail.

4.5. Geographic patterns

Recent literature suggests a possible role for geographic differences in explaining the mortality growth of the less educated (Montez et al., 2019a,b). 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 Fig. B11).

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 Fig. B12). Specifically, census divisions with larger Q_4 shares had greater percentage reductions in death rates, while the reverse is true for areas with larger population shares of Q_2. 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 Q_1, rather than positive as it is for Q_2, which argues against such a story. Taken together, this analysis provides little support for geography-based explanations for our main results.

5. Discussion

We provide a detailed analysis of mortality changes, from 2001 to 2017, for subgroups stratified by sex, race, 5-year age groups and education quartiles. Consistent with prior work, we find the least-educated women generally experienced the worst mortality trends. However, this is less true for males, where all but the highest education quartile had substantially similar patterns. Conversely, both males and females in the top quartile had the most favorable mortality experiences, on average, leading to a widening gap between

[11] Although our analysis uses single years of education to construct quartiles, we calculate the inverse coefficients of variation 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.

them and the rest of the population. An important point, however, is that there is substantial variation around these averages. Within sex-race-age groups, monotonicity violations are frequent—where lower education quartiles have more favorable mortality trends 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. These non-monotonic patterns are particularly common for males, but also occur reasonably often for females, and also for nonwhites of both sexes.

Our results should be interpreted in light of several limitations. First, causation is generally not possible 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 heterogeneity analysis estimates linear trends to concisely summarize mortality patterns across groups, although non-linearities could sometimes exist. This is unlikely to be a major issue since our preceding estimates of average effects show qualitatively similar results using alternative approaches that either assume linearity over shorter periods or only consider the change between the starting and ending sample years. Fourth, we have 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, immigration patterns are probably not driving our findings. Finally, our methods account for rising levels of schooling, but do not consider 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, suggesting that within-quartile changes in educational attainment are unlikely to meaningfully affect the interpretation of our results.

We view it as advantageous to evaluate both relative and absolute changes in death rates. 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 40–54 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. Also, as already discussed, log death rates are likely to be more useful when aggregating the effects of varying groups, since using levels is likely to overweight groups with high baseline death rates.

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 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 et al., 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.

Better understanding the trends in and, ultimately, the determinants of mortality is especially important given the relatively poor U.S. performance in improving life expectancy in recent years. For instance, a recent study finds that 15–64 year olds in the U.S. have some of the worst mortality trends among 16 peer countries examined, with particularly poor outcomes recently emerging among young adults (National Academies of Sciences, Engineering, and Medicine, 2021). Within the U.S., recent work suggests that differences in state policies, across a variety of domains, are associated with changes in life expectancy and any causal effects are likely to differ across education groups (Montez et al., 2020). The linkages between education and mortality are assuredly complex. While numerous factors determine the risk of death, education is likely to be critical given its influence on many of them – such as access and adherence to medical advice, availability and consumption of high quality food, and health behaviors such as physical activity and disease prevention. However, this study also clearly demonstrates the risks of making overly broad generalizations about education patterns and suggests the importance of future work that extends analyses of this type to consider specific causes of death.

Authors' statement

Adam Leive and Christopher Ruhm were equally involved in all aspects of this paper including formulation of the ideas and methods, descriptive and econometric estimates and drafting of the manuscript.

Supplementary materials

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.jhealeco.2021.102494.

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