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Trends in the Educational Gradient of U.S. Adult Mortality from 1986 to 2006 by Race, Gender, and Age Group*

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Abstract

The educational gradient of U.S. adult mortality became steeper between 1960 and the mid 1980s, but whether it continued to steepen is less clear given a dearth of attention to these trends since that time. This study provides new evidence on trends in the education-mortality gradient from 1986 to 2006 by race, gender, and age among non-Hispanic whites and blacks using data from the 2010 release of the National Health Interview Survey Linked Mortality File. Results show that, for white and black men, the gradient steepened among older ages because declines in mortality risk across education levels were greater among the higher educated. The gradient steepened among white women, and to a much lesser and only marginally significant extent among black women, largely because mortality risk decreased among the college-educated but increased among women with less than a high school degree. Greater returns to higher education and compositional changes within educational strata likely contributed to the trends.

Keywords

education; mortality; educational gradient; NHIS-LMF

The inverse association between socioeconomic status and adult mortality risk in the United States is a social fact. After the seminal work of Kitagawa and Hauser (1973), numerous and more recent studies have demonstrated that adult mortality risk is inversely associated with socioeconomic status whether measured by educational attainment, income, or occupational characteristics (Elo and Preston 1996; Hummer and Lariscy 2010; Jemal et al. 2008a; Molla, Madans, and Wagener 2004; Moore and Hayward 1990; Preston and Taubman 1994; Rogers, Hummer, and Nam 2000; Rogot, Sorlie, and Johnson 1992). Research in the United States has increasingly focused on educational attainment largely because, compared with income and occupation, it is a more stable measure of socioeconomic status, it is available for men, women, and individuals outside of the labor force, and it is most closely associated

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with health behaviors (Winkleby et al. 1992). During the decades following Kitagawa and Hauser's seminal work, comparisons of the inverse association (i.e., the "educational gradient") between 1960 and the mid 1980s revealed another intriguing pattern: although mortality rates declined for the U.S. population overall, they declined more rapidly for higher-educated adults, creating a much steeper educational gradient (Crimmins and Saito 2001; Feldman et al. 1989; Lauderdale 2001; Pappas et al. 1993; Preston and Elo 1995; Rogot, Sorlie, and Johnson 1992).

Whether the educational gradient of mortality stabilized, flattened, or steepened since the mid 1980s is less clear, although emerging evidence suggests that it continued to steepen (Cutler et al. 2010; Jemal et al. 2008; Meara, Richards, and Cutler 2008). Here, we provide new evidence on recent trends in the educational gradient of U.S. adult mortality from 1986 to 2006. We address the following questions. Did the gradient stabilize, flatten, or steepen from 1986 to 2006? If the gradient changed, for which education levels and for which age groups did it change? We address these questions separately for non-Hispanic white and black men and women. We use the 2010 release of the public-use National Health Interview Survey Linked Mortality File (NHIS-LMF), which links adults from the 1986 through 2004 annual waves of the National Health Interview Survey with death records in the National Death Index through December 31, 2006. The NHIS-LMF contains more than 1.3 million adults experiencing 177,597 deaths during the mortality follow-up period. In the following sections, we first discuss key structural trends that may have influenced the strength of the educational gradient in recent decades, and we then review historical trends in the gradient.

STRUCTURAL TRENDS

The United States witnessed an unprecedented increase in life expectancy over the twentieth century. Life expectancy at birth, for example, increased from 47 to 78 years between 1900 and 2006 (Arias 2010). The pace of increase fluctuated by race and gender throughout the century and, at least until the mid 1980s, the pace was more rapid for higher educated individuals (Crimmins and Saito 2001; Feldman et al. 1989; Lauderdale 2001; Pappas et al. 1993; Rogot et al. 1992). Whether the pace continued to be more rapid for higher-educated individuals since the 1980s is less clear given a dearth of attention to these trends. We can offer several hypotheses, however. On one hand, the educational gradient of adult mortality may have flattened due to certain structural trends during the close of the twentieth century. For example, health promotion goals were elevated on the national agenda. Healthy People 2000 and 2010 progressively sought to reduce and ultimately eliminate heath disparities. Cautiously optimistic, the Healthy People 2000 progress report concluded that 58 percent of targeted health disparities between low socioeconomic groups and the population overall had been reduced or eliminated (National Center for Health Statistics 2001). Furthermore, the gradient may have flattened if the mortality rates of highly educated individuals have encroached upon the probabilistic upper limits of human longevity (Carnes and Olshansky 2007; Fries 1980), allowing the mortality rates of the less educated to gradually converge.

However, other structural trends portend a steepening educational gradient. Several of these trends imply that the gradient steepened because education "buys more" today, particularly with respect to labor market opportunities and resources for creating a healthy lifestyle. For example, the transition from a diversified economy with good jobs in both blue and white collar sectors to an hourglass economy means that good jobs are increasingly limited to the high-tech, professional sector and require college educations (Portes and Zhou 1993). Thus, jobs with high incomes, health care benefits, and retirement security—resources that enhance health—are increasingly restricted to the most highly educated individuals. Another trend entails innovations in medical technology. Paradoxically, these innovations can widen socioeconomic disparities in health because individuals with greater access to resources such as knowledge and money also have greater access to these innovations (Frisbie et al. 2004; Glied and Lleras-Muney 2008; Phelan et al. 2004). A third trend concerns health behaviors, particularly smoking patterns. Prior to the 1964 Surgeon General's Report, smoking patterns were fairly similar across education levels. Since then, higher-educated individuals are less likely to smoke and are more likely to quit smoking than individuals with less education. In 2007 and among individuals aged 18 and over, the percentage of current smokers was 30 percent among those with less than a high school diploma, 13 percent among those with an undergraduate degree, and just 6 percent among those with a graduate degree (Centers for Disease Control 2008). In general, higher-educated individuals more rapidly integrate new health-related knowledge into their lifestyle. A fourth, potentially important trend concerns the quantity and quality of social ties such as marriage. The better educated have become even more likely than their less-educated peers to get and stay married (DiPrete and Buchmann 2006); and trends toward greater educational homogamy among spouses (Schwartz and Mare 2005) and nonfamily ties (McPherson, Smith-Lovin, and Brashears 2006) suggest that the better educated may now yield even greater health returns from social networks.

The trends above are well known and often invoked to explain the growing health and mortality advantage of higher-educated adults. Often overlooked, however, are compositional changes *within* education strata over time. For example, as the proportion of adults with less than a high school diploma declined, this lower-educated group may have become more homogeneous and disadvantaged, whether disadvantage stems from low parental socioeconomic status, stressful childhood home environments, poor academic performance, behavioral problems, or other forms of adversity (Alexander, Entwisle, and Kabbani 2001; Rumberger 1995). At the same time, the secular rise in education levels suggests that higher education groups may have become more heterogeneous and less select. In summary then, the educational gradient of mortality may have flattened since the mid 1980s due to policy initiatives, current limits to human longevity, and compositional changes that increased heterogeneity among the higher-educated. Alternatively, the gradient may have steepened if, for example, higher education buys more today in terms of labor market opportunities and ability to create a healthy lifestyle, and if compositional changes increased homogeneity and disadvantage among the lower-educated.

HISTORICAL TRENDS IN THE GRADIENT

The bulk of evidence in this literature compares the educational gradient in 1960 to the mid 1980s among white men. As early as the mid 1970s, the gradient among white men began to steepen (Duleep 1989) and it continued to do so throughout the 1980s (Crimmins and Saito 2001; Feldman et al. 1989; Lauderdale 2001; Pappas et al. 1993; Preston and Elo 1995; Rogot et al. 1992). Between 1970 and 1990, the gap in life expectancy at age 30 between white men with a primary education (0–8 years) and those with more than a high school diploma grew from 4.1 to 6.7 years (Crimmins and Saito 2001). Although absolute mortality rates declined across education levels during this period, they declined faster for higher-educated men largely due to their disproportionate declines in heart disease mortality (Feldman et al. 1989), which correlates with their sharper reduction in smoking.

In contrast to the consistently reported increase in the gradient among white men, reported trends among white women during that timeframe show little consistency. Rogot and colleagues (1992) found a flattening gradient among working-age and older white women, in contrast to the steepening gradient they found among white men. Others concluded that mortality declines occurred fairly evenly across education levels among white women, so their gradient was largely unaffected (Feldman et al. 1989). Still others reported steepening gradients (Lauderdale 2001; Pappas et al. 1993), or cohort-specific trends (Crimmins and

Saito 2001; Preston and Elo 1995). Cohort differences may be particularly salient for women, given dramatic shifts in labor market attachment and family formation during that time. Indeed, Crimmins and Saito (2001) found that the gradient steepened for white women under 54 years, yet flattened for women 55 years and older from 1960 to 1990. Overall, however, the gap in life expectancy at age 30 between white women with a primary education and those with more than a high school education *decreased* from 4.7 to 3.8 years from 1970 to 1990 (Crimmins and Saito 2001).

Comparatively few studies have examined trends in the gradient among black men and women. Historically, the gradient appeared steeper for black men than white men in part because low educational attainment was particularly detrimental for black men (Crimmins and Saito 2001; Lin et al. 2003), although this difference may no longer exist (Cutler and Lleras-Muney 2006). Between 1960 and the 1980s, the gradient steepened among black men (Crimmins and Saito 2001; Pappas et al. 1993). For example, the gap in life expectancy at age 30 between black men with a primary education and those with more than a high school diploma grew from 5.4 to 11.8 years from 1970 to 1990 (Crimmins and Saito 2001). Historically, the gradient also appeared steeper for black women compared with white women (Crimmins and Saito 2001; Lin et al. 2003), although this difference also may no longer exist (Cutler and Lleras-Muney 2006). Between 1960 and the 1980s, the gradient steepened for black women, and the steepening appeared sharper for black women than white women (Crimmins and Saito 2001; Pappas et al. 1993). The gap in life expectancy at age 30 between black women with a primary education and those with more than a high school diploma grew from 5.7 to 10.5 years from 1970 to 1990 (Crimmins and Saito 2001).

Emerging evidence suggests that the educational gradient of adult mortality continued to steepen during the 1990s (Cutler et al. 2010; Jemal et al. 2008; Meara et al. 2008). Cutler and colleagues used various waves of the National Health and Nutrition Examination Surveys (NHANES) and an earlier version of the NHIS-LMF that we use here to examine the gradient between the early 1970s and early 2000s for non-Hispanic whites 25–74 years of age. Using a dichotomous indicator of education (0–12 years, 13+ years), they found that the education gap in mortality risk widened for men and especially for women. Using vital statistics data, Jemal et al. (2008b) and Meara et al. (2008) reported that the gradient steepened among non-Hispanic whites and blacks during the 1990s. Meara and colleagues also replicated their analysis for non-Hispanic whites using the National Longitudinal Mortality Study (NLMS). Although both studies concluded that the gradient steepened during the 1990s, a few discrepancies are noteworthy. To begin, Meara and colleagues reported that the steepening was more pronounced among women than men, while the results from Jemal and colleagues suggested the opposite (Jemal et al. 2008: see table 1). In addition, Meara found that the steepening among black women was large and statistically significant, while Jemal found it was modest and not significant. Differences in analytical samples may explain the discrepancies; these include differences in ages (Jemal et al. analyzed ages 25–64; Meara et al. analyzed ages 25–84), time periods (Jemal et al. compared 1993 to 2001; Meara et al. compared 1990 to 2000, and compared 1981–85 to 1991–95 using the NLMS for whites), and education categories (Jemal et al. included 0–11, 12, 13–15, and 16+ years; Meara et al. included 0–12 and 13+ years).

Our research builds on these informative analyses in a number of ways. We use the 2010 release of the public-use NHIS-LMF to analyze trends in the educational gradient of U.S. adult mortality from 1986 through 2006 for non-Hispanic white and black men and women. This version of the NHIS-LMF contains the most current survey data for examining the gradient. A major advantage of this data is that it contains respondent-provided educational attainment. In contrast, educational attainment in the vital statistics data are provided by a third party at the time of death, with inaccuracies that vary by race and age of the decedent

(Sorlie and Johnson 1996); thus, analyses using vital statistics data often dichotomize education to mitigate potential inaccuracies (e.g., Meara et al. 2008). The detailed data on educational attainment within the NHIS-LMF allow us to examine a four-category measure of education, thereby providing a finer assessment of mortality trends across education levels. Another advantage of the NHIS-LMF is its individual matching of survey records to death certificates, which avoids the numerator-denominator mismatch that often occurs when using the ratio of death counts to census counts for mortality estimates. Another strength of this data is its large sample size, which allows us to stratify our analyses by race-gender-age groups. Stratification is important because: (1) the strength of the education-mortality association may differ for men versus women (Elo and Preston 1996; Montez et al. 2009), for whites versus blacks (Crimmins and Saito 2001; Lin et al. 2003), and for younger versus older adults (Lauderdale 2001; Lynch 2003) during any given period of time because the resources that education confers may vary by race, gender, and age, and (2) over time, the strength of the association may change at different paces for different race-gender-age groups. For instance, the returns from higher education for one's economic standard of living (DiPrete and Buchmann 2006), probability of getting and staying married (DiPrete and Buchmann 2006), and smoking behavior (Cutler et al. 2010; Meara et al. 2008)—potentially important mediators of the education-mortality association—grew during the last four decades of the twentieth century, but by varying degrees for race-gender-age groups. Given that the mechanisms underlying trends in the education-mortality association are highly complex and not completely understood (see Cutler et al. 2010), we did not develop detailed hypotheses about which groups experienced greater or lesser changes in the gradient during our study period. However, based on recent studies of the gradient (Cutler et al. 2010; Meara et al. 2008), we anticipated that the gradient steepened more among women than men, particularly among whites.

DATA AND METHODS

Data

We used the 2010 release of the public-use National Health Interview Survey Linked Mortality File (NHIS-LMF), which links adults in the 1986 to 2004 annual waves of the National Health Interview Survey (NHIS) with death records in the National Death Index (NDI) through December 31, 2006. The NHIS is a cross-sectional, annual household interview that has been conducted since 1957. It is the principal source of health information on the civilian, non-institutionalized population of the United States (National Center for Health Statistics 2005). Indeed, Preston and Taubman (1994:291) described the NHIS as "the most authoritative source of national data on socioeconomic differences in health status." The NDI is a computerized database of all certified deaths in the United States since 1979. The 2010 NHIS-LMF links adults in the NHIS to death records in the NDI using a probabilistic matching algorithm (Lochner et al. 2008; National Center for Health Statistics 2009). The algorithm correctly classifies the vital status of 98.5 percent of eligible survey records, although the probability of correct matches is marginally lower for older ages, white females, and extended mortality follow-up durations (Ingram, Lochner, and Cox 2008). The 2010 NHIS-LMF also utilizes vital status information from the Social Security Administration (SSA) and Centers for Medicare and Medicaid Services (CMS). Adults identified as deceased in any of these sources (NDI, SSA, CMS) are identified as decedents in the NHIS-LMF. The public-use 2010 NHIS-LMF contains 1,326,350 match-eligible adults at least 18 years of age at the time of NHIS interview, with 177,597 (13.4 percent) of these adults identified as subsequent deaths by December 31, 2006.

Sample

Our analytic sample consists of a person-year file that includes non-Hispanic whites and non-Hispanic blacks 45 to 84 years of age during 1986 to 2006. We created this analytic sample in two main steps. In the first step, we generated a person-year file in which we aged the match-eligible adults 18 years and older by one year beginning with their interview year until their year of death, or until 2006 if they survived the follow-up period. For the 0.5 percent of adults who were missing information on either month or year of birth, we imputed month of birth by random assignment and we imputed year of birth by subtracting their age from their interview year.

In the second step, we selected our demographic sample of interest from the person-year file above. We first selected non-Hispanic whites and non-Hispanic blacks. We excluded other race/ethnic groups because of the limited number of deaths for other groups, the greater potential for education to have been obtained abroad among groups with high levels of immigration, and because preliminary analyses showed less certainty in death matching for other groups. Among the non-Hispanic whites and blacks, we then selected adults who met two age criterion. Our first criterion was that the adult was 25 to 84 years of age at the time of their NHIS interview. We set the lower limit at 25 because educational attainment is often incomplete before then, while we set the upper limit at 84 because the NHIS began top coding age at 85 in 1997. Our second criterion was that the age of the person-year record be between 45 and 84. We set the lower limit at 45 due to the small number of deaths below age 45, while we set the upper limit at 84 because preliminary analyses indicated that the matching of death certificates among women after roughly 85 was less successful than it was for men, which could influence our results. Given these two criterion, adults could "age in" or "age out" of our person-year file. As an example of "aging in," adults 35 years of age during their 1990 NHIS interview—while meeting our first age criterion—did not contribute person-year records to our data file until they reached 45 years of age in 2000, assuming they survived to 2000. As an example of "aging out," adults 80 years of age during their 1990 interview year stopped contributing person-year records to our data file once they reached 85 years of age in 1995. Lastly, from these person-year records we selected adults who provided their educational attainment (0.8 percent missing). Our final analytic sample contained 7,144,521 person-year records and 111,574 deaths. Note that we did not delete person-year records that did not meet our criteria because this could cause the standard errors from our models to be computed incorrectly. Rather, we retained all person-year records from match-eligible adults 18 years and older, and simply distinguished records in our final analytic sample from records in the remainder of the sample using a dichotomous indicator, which we then used in SUDAAN as recommended (National Center for Health Statistics 2005).

Methods

We estimated Poisson regression models to evaluate whether the educational gradient of mortality changed during the 1986 to 2006 period for each defined race-gender-age group. The Poisson models estimated the natural logarithm of the annual mortality rate as a linear function of our covariates, and they accounted for the smaller exposure interval for person-year records tied to respondents' NHIS interview year and their year of death. All models were estimated with SUDAAN (SUDAAN 2005), weighted using the eligibility-adjusted sample weights, and adjusted for the complex sample design of the NHIS-LMF.

Our covariates included age, education, trend, and the interaction between education and trend. Age is a time-varying, continuous variable ranging from 45 to 84 years. Education was included as four categories (three dummy variables) indicating adults with less than a high school diploma, a high school diploma, some college, or a bachelor's degree or higher (omitted reference). The category "some college" includes adults who attended a four-year university but did not graduate with a bachelor's degree, as well as adults with a technical, vocational, or academic associate's degree. We did not distinguish 0 to 8 years of education due to the small sample size of this group in our data and the U.S. population. Trend is an ordinal variable indicating three 7-year periods within the 21 years of follow-up, 1986 to 1992 (trend=1), 1993 to 1999 (trend=2), and 2000 to 2006 (trend=3). We selected this specification because we found in preliminary analyses comparing various specifications of calendar years (e.g., single years, 7 groups of 3 years) that it sufficiently smoothed the annual variation in mortality risk within the NHIS-LMF, and would therefore provide stable model estimates. To be sure, we conducted sensitivity analyses using two alternative specifications of calendar years. One specification was a categorical measure of the three 7-year periods to test for nonlinear trends in the gradient (preliminary analyses did not indicate nonlinearities, however), and was included in the models as two dichotomous indicators (1986 to 1992 omitted, 1993 to 1999, 2000 to 2006). The second specification was a linear measure which included all single years of follow-up except for 1986, which exhibited a particularly high degree of variability. We report the sensitivity results below.

We stratified our models by race-gender-age group. Age groups include the entire 45–84 year age range, as well as 10-year subgroups for ages 45–54, 55–64, 65–74, and 75–84 to assess whether educational differences in mortality changed within certain or all portions of the adult life course. For each race-gender-age group, we estimated one model that included age, education, and trend, and a second model that added the education-by-trend interactions. We also conducted likelihood ratio tests to determine whether the education-by-trend interactions significantly improved the fit of the models for each race-gender-age group.

RESULTS

Tables 1 and 2 contain distributions of educational attainment within each 7-year period by race, gender, and age. The distributions reflect person-year records to be consistent with our models; however, for ease, we discuss the distributions in terms of individuals. The tables illustrate the secular rise in educational attainment for cohorts born in the twentieth century. During the first period (1986 to 1992), almost one-half of whites (47.1 percent of men and 46.3 percent of women) and more than three-fourths of blacks (78.4 percent of men and 75.9 percent of women) in the oldest age group did not have a high school diploma. By the third period (2000 to 2006), just 8.8 percent of white men, 7.7 percent of white women, 18.0 percent of black men, and 17.2 percent of black women in the youngest age group did not have a high school diploma. Similarly impressive trends in college completion also occurred during the study period.

Table 3 presents coefficients from Poisson regression models that formally test for trends in the educational gradient of mortality among white adults. We first discuss the results for white men. Model M2 tests for trends in the gradient among white men aged 45–84, and shows that all trend-by-education interaction coefficients are statistically significant and positive, indicating that the gradient became steeper for white men during the 1986 to 2006 period. Also supporting this conclusion, the interaction terms significantly improve the model fit compared with the main effects model ($\chi^2 = 12$, $p<0.01$). In the next four columns, the age-stratified models reveal that the steeper gradient among the overall group of men aged 45–84 was primarily due to a steeper gradient among men aged 65–84. Among men aged 65–74, all four education levels experienced declines in mortality risk during our study period; however, the declines were greater for higher education levels (e.g., the slope was greatest for college-educated men (-0.192) and smallest for men without a high school diploma (-0.192 + 0.125 = -0.067)). Similar to men aged 65–74, among men aged 75–84 all four education levels experienced declines in mortality risk over the study period, although declines among men with less than a high school education and those with some college did not keep pace with declines among men with a college degree, thereby creating a steeper education-mortality gradient. See Figure 1 for an illustration of these trends.

Before we discuss the results for white women, a comment about the positive coefficient for trend in model W1 in Table 3 is needed. This model shows an increase in mortality risk, net of age and the main effect of education, during the 1986 to 2006 period. Ancillary analyses and the remaining models in Table 3 suggest that the positive coefficient is due to: (1) an increase in mortality risk among white women aged 75–84 in the NHIS-LMF, which may reflect trends in nursing home institutionalization, and (2) an increase in mortality risk among low-educated women in all four age groups. Regarding point (1), according to 1990 and 2000 vital statistics data, death rates for all four age groups declined over the decade, albeit by a modest 1.4 percent for the 75–84 group (U.S. Census Bureau 2008). Trends in nursing home use may underlie the discrepancy between the vital statistics and the NHIS-LMF. Between 1985 and 1997, the rate of nursing home use declined (Sahyoun et al. 2001). It declined in part because disability rates declined, but largely because of emerging, non-institutional alternatives to long-term care, such as home health care, assisted living, community-based centers, and continuing-care retirement communities (Bishop 1999; Gallagher 2000). Thus, these least healthy adults who would have previously been excluded from the NHIS-LMF non-institutional sampling frame were now eligible to participate, which could inflate mortality risks during our study period. Because nursing home residents are mostly white women (Sahyoun et al. 2001), these trends exert the largest impact on their results.

Model W2 in Table 3 tests for trends among white women aged 45–84, and shows that all trend-by-education interaction coefficients are statistically significant and positive, indicating that the gradient became steeper for white women during the 1986 to 2006 period. Also supporting this conclusion, the interaction terms significantly improve the model fit ($\chi^2 = 41$, $p<0.01$). The age-stratified models reveal that the gradient became steeper for all four age groups, with the most pronounced changes experienced by the youngest group. In contrast to white men whose steepening gradient was due to differential declines in mortality risk across education levels, the steepening gradient among white women was largely due to declines in mortality risk among college-educated women (except for women aged 75–84) alongside *increases* in mortality risk among women with less than a high school diploma across all four age groups. Clearly then, the 1986 to 2006 period exhibited substantial changes in the education-mortality association among young white women in particular, with low-educated women losing statistically significant ground against their college-educated peers. Even women aged 45–54 with a high school diploma experienced an increase in mortality risk across the period. Figure 2 illustrates the trends in the gradient among white women.

Table 4 contains coefficients from Poisson regression models for black adults. Model M2 tests for trends among black men aged 45–84, and shows that all trend-by-education interaction coefficients are statistically significant and positive, indicating that the gradient became steeper for black men during the 1986 to 2006 period; however, two of the interaction terms are only marginally significant and the inclusion of all three terms did not significantly improve the model fit ($\chi^2 = 4$, $p>0.10$). The age-stratified models reveal that the steeper gradient among the overall group of men aged 45–84 was primarily due to a steeper gradient among men aged 65–74, and that the interaction terms significantly improved that age-stratified model ($\chi^2 = 8$, $p<0.05$). Table 4 and Figure 3 illustrate that the steepening gradient among black men aged 65–74 was mainly due to greater declines in mortality risk among college-educated men (slope = -0.351) compared with less-educated men: men with less than a high school diploma experienced a much smaller decline (slope = -0.351 + 0.296 = -0.055) as did men with some college (slope = -0.093), while men with a high school diploma experienced an increase (slope = 0.045).

Similar to white women, the trend coefficient for black women in Model W1 within Table 4 is positive (but not statistically significant), which may reflect increases in mortality risk among low-educated black women and, to a lesser extent than for white women, trends in nursing home institutionalization among elderly black women. Model W2 tests for trends among black women aged 45–84 and shows that just one interaction coefficient is marginally significant, and that the interaction terms significantly improve the model fit ($\chi^2 = 9$, $p<0.05$). Thus, the gradient became marginally steeper for black women during the study period. The age-stratified models found statistically significant changes in the gradient only among the 75–84 group, but these interactions did not improve the model ($\chi^2 =4$, $p>0.10$). Thus, the results for black women are somewhat inconclusive. However, taken together, they directionally indicate that the gradient became marginally steeper among black women aged 45–84 because women with less than a high school diploma may have experienced an increase in mortality risk across all four age groups, while college-educated women may have experienced a decrease in mortality risk across three age groups (except possibly the 45–54 group), and that the most pronounced changes in the gradient may have occurred among elderly black women.

We also conducted sensitivity analyses using different specifications of time across the study period and then replicating Tables 3 and 4. Recall that one specification was a categorical measure of the three, 7-year periods which we included in the models as two dichotomous indicators (1986 to 1992 omitted, 1993 to 1999, 2000 to 2006). We found that models using the ordinal measure were preferred over models using the categorical measure based on the Bayesian Information Criterion. With that important caveat in mind, the categorical measure identified changes in the gradient for the same age groups of white men and women as the ordinal measure, but they did not identify changes among black adults. Also, the expansion of the gradient among whites was most pronounced in the third period with one exception (white women aged 75–84). The second specification was a linear measure containing single years of follow-up excluding 1986. In general, these results corroborated those using the ordinal measure. In the few cases that the results differed, the differences were not practically meaningful and they appeared to derive from less stable estimates using the linear measure due to annual variation in the data.

DISCUSSION

In summary, our results revealed that the educational gradient of mortality among U.S. adults aged 45–84 became steeper across the 1986 to 2006 period for some race-gender-age subgroups. In other words, the gap in mortality risk between lower- and higher-educated adults in these subgroups expanded to create even larger disparities in the length of life among many Americans—continuing a trend that began at least as early as the mid twentieth century—despite major policy initiatives designed to reduce socioeconomic disparities in health. The expanding educational gap in mortality risk does not imply that those policy initiatives were ineffective: they simply imply that the gains made by these initiatives in reducing mortality disparities were overshadowed by stronger structural forces working against them (see also Meara et al. 2008).

The existence and magnitude of the expansion in mortality risk across education levels varied by race-gender-age group. Among white men, the steepening gradient was concentrated among older age groups (65–84 years) and resulted from declines in mortality risk among the college-educated that outpaced declines among their less-educated peers. In contrast, white women across the full 45–84 year age range experienced a steeper education-mortality gradient, primarily due to decreases in mortality risk among college-educated women alongside *increases* in mortality risk among women with less than a high school diploma. In fact, young (aged 45–54) white women exhibited the most pronounced expansion of the gradient across all race-gender-age groups, and experienced an increase in mortality risk for those up to and including a high school diploma. Similar to white men, the expansion of the gradient among black men was concentrated among older ages (65–74 for black men compared with 65–84 for white men) and largely resulted from greater declines in mortality risk among the college-educated that outpaced declines among the less-educated. Lastly, among black women, the gradient became marginally steeper during the study period although our results did not unequivocally identify which age groups experienced these changes. However, the results directionally indicated that the gradient became marginally steeper because—similar to white women—black women with less than a high school diploma may have experienced an increase in mortality risk alongside a decrease in risk among the college-educated (except possibly ages 45–54). Of course, given the cross-sectional nature of the NHIS-LMF, all of our results are contingent upon individuals surviving to at least age 45.

Increasing returns to higher education (whether economic, social, behavioral, or otherwise), as well as compositional changes within educational strata, may have contributed to the steepening education-mortality gradient. For instance, regarding economic returns to education, DiPrete and Buchman (2006) compared the standard-of-living returns to education for young (25–34 years) white and black males and females with a college degree against their peers with a high school diploma across the 1964 to 2002 Current Population Surveys and found that the standard-of-living gap grew over time, among women more than men, and among whites more than blacks—patterns that roughly correspond with changes in the education-mortality gradient reported here.

Similarly, the returns to higher education may have grown with respect to social ties. For instance, the returns to education have increased for the probability of getting and staying married, among women more than men, and among whites more than blacks (DiPrete and Buchmann 2006)—again, patterns that correspond with changes in the gradient reported here. These marriage trends could have implications for the education-mortality gradient because married adults exhibit lower mortality risk than the unmarried (Lillard and Waite 1995), in part through more favorable economic circumstances, health behaviors, and social support (Lillard and Waite 1995; Umberson 1987; Waite 1995). Higher-educated adults may not only be more likely to get and stay married than their less-educated peers in recent decades, they may also yield greater returns from marriage given the increasingly central role of education in assortative mating (Schwartz and Mare 2005), along with evidence that the education of both spouses contributes to each partner's mortality risk (Kravdal 2008; Montez et al. 2009). Educational homogamy of nonfamily social ties has also increased (McPherson et al. 2006), which could strengthen the education-mortality gradient because social networks influence a host of health-related behaviors and outcomes (see Smith and Christakis 2008). Taken together, trends in the quantity and quality of social ties may have contributed to trends in the education-mortality gradient, especially considering that social relationships, or the lack thereof, represent a risk factor for health which rivals well-established behavioral risk factors such as smoking (House, Landis, and Umberson 1988).

The returns to higher education have also increased with respect to health behaviors. For example, across U.S. birth cohorts that correspond to the ages of adults in our data, smoking prevalence among blacks and whites declined faster for higher-educated adults (those with a high school diploma or more) and for men than women, while prevalence actually increased among low-educated women (Escobedo and Peddicord 1996). Other indicators of a healthy lifestyle, such as physical activity, obesity, and having a usual place of health care also exhibit substantial gaps across education levels (Pleis and Lethbridge-Çejku 2007). However, the extent to which behavioral factors have contributed to the steepening education-mortality gradient over time is unclear. Cutler and colleagues (2010) examined whether obesity and smoking trends explained the expanding education-mortality gradient among whites aged 25–74 between the early 1970s and early 2000s, but found little supporting evidence. The strongest evidence implicated smoking trends among women, but even so, smoking explained less than 10 percent of the steeper education-mortality gradient in their data. Yet, another study that decomposed trends in the overall education-mortality gradient by cause of death found that two causes that share smoking as a major risk factor (lung cancer and chronic obstructive pulmonary diseases) explained more than 20 percent of the steeper gradient between 1990 and 2000 among an aggregated sample of whites and blacks aged 25–84, and roughly 25 percent among white women aged 45–84 (Meara et al. 2008).

While increasing returns to higher education may have pulled the gap wider from the top of the education distribution, compositional changes may have simultaneously pulled the gap wider from the bottom. Indeed, the confluence of child labor and compulsory education laws in the early twentieth century and an increasingly urban and industrialized economy have redefined what it means to terminate formal education before high school completion: these individuals have likely become a more homogeneous and disadvantaged group. Such compositional changes could help explain why women without a high school diploma experienced an increase in mortality risk in our study (see also Meara et al. 2008): low-educated women may have become more select than their male peers given that women became less likely than men to drop out of high school starting around 1980 and more likely to graduate college starting in the mid 1980s (Freeman 2004). Compositional changes within the top end of the education distribution could also have steepened the gradient if college-educated adults became more positively select. However, the composition of college graduates probably became less select among the adults in our study given both the rise in college enrollment across income levels as well as the G.I. Bill of Rights, the latter of which opened doors to higher education to a wider group of individuals than ever before.

Many other factors have likely contributed to the steeper education gap in mortality risk (Cutler et al. 2010). These factors likely vary by race-gender-age, and across time because both the type and strength of mechanisms that persistently link socioeconomic status with mortality vary across time (Phelan and Link 2005). The point here is that it is unlikely that there exists a single, silver bullet explanation for the expanding education-mortality gradient that we documented.

In conclusion, our study showed that the gap in U.S. mortality risk between lower- and higher-educated adults expanded between 1986 and 2006 among certain demographic subgroups to create even larger disparities in the length of life among many Americans, and that the expansion was particularly pronounced among young white women. The expansion occurred despite policy initiatives to reduce health disparities among Americans. As we stated earlier, our results do not imply that these initiatives were unsuccessful—only that the structural forces working against them were stronger. But, what are these structural forces? We speculated on three of them here—economics, social ties, and smoking—but we encourage future research to identify the structural forces by systematically examining a more comprehensive set of potential mechanisms than have been examined so far. Indeed, the multifarious mechanisms linking education with health and mortality at any point in time, and over time and across birth cohorts, extend beyond the usual suspects, income and smoking (e.g., Ross and Wu 1995). Other potential mechanisms include changes in social ties, neighborhood living environments, access to medical care, job fulfillment, occupational hazards, psychosocial resources, and availability and sophistication of health-related information, for example (see Cutler et al. 2010). Insights from research on mechanisms such as these would be invaluable for policy initiatives by directing them to the most efficacious areas for investment. Finally, future research should also continue to monitor trends in the education-mortality gradient because, as history has shown, it is dynamic and may change as the mechanisms that link education and mortality shift over time (Phelan and Link 2005).

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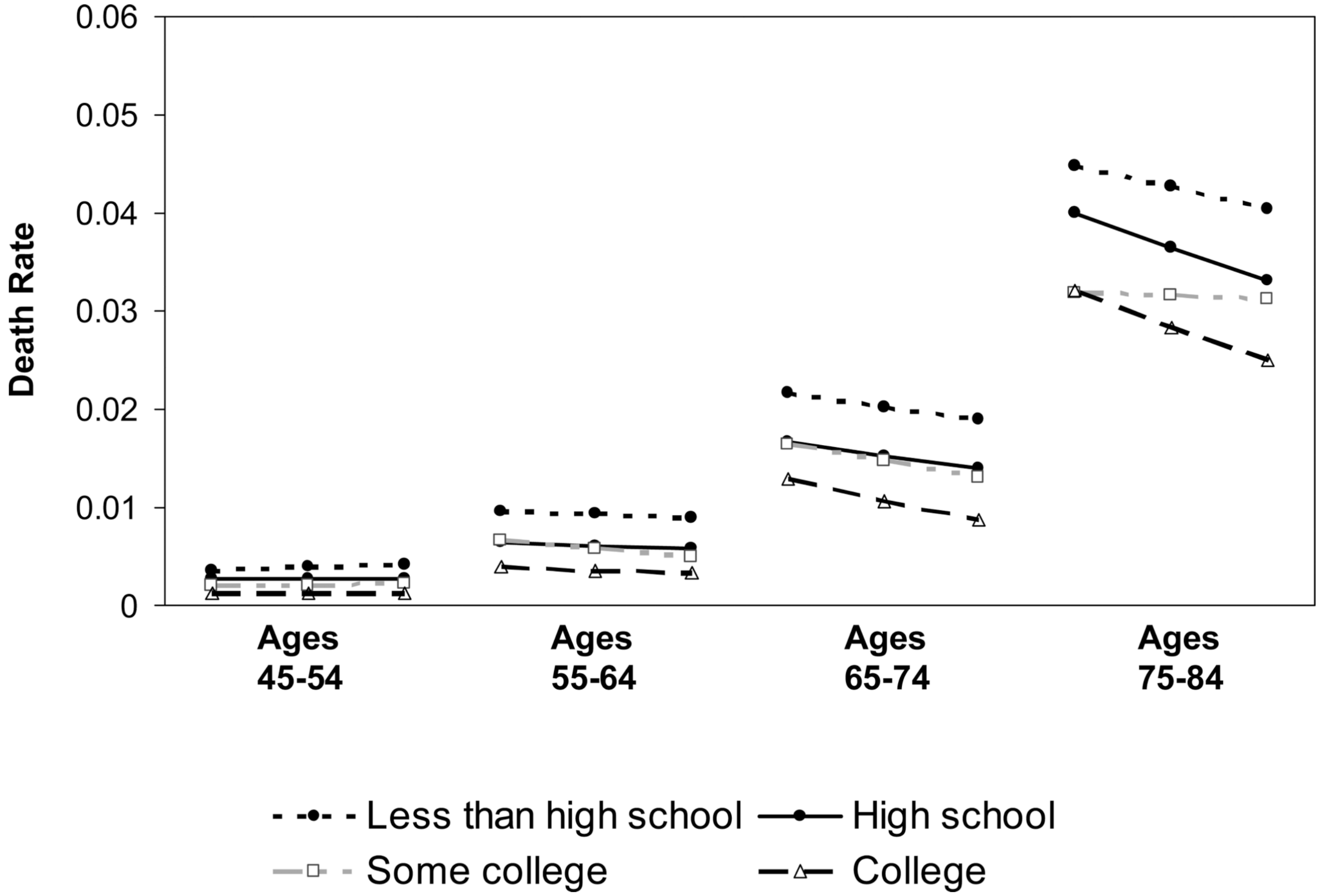

Figure 1.
Trends in Annual Death Rate for Non-Hispanic White Men across Three Time Periods (1986–1992, 1993–1999, 2000–2006) by Educational Attainment and Age Group

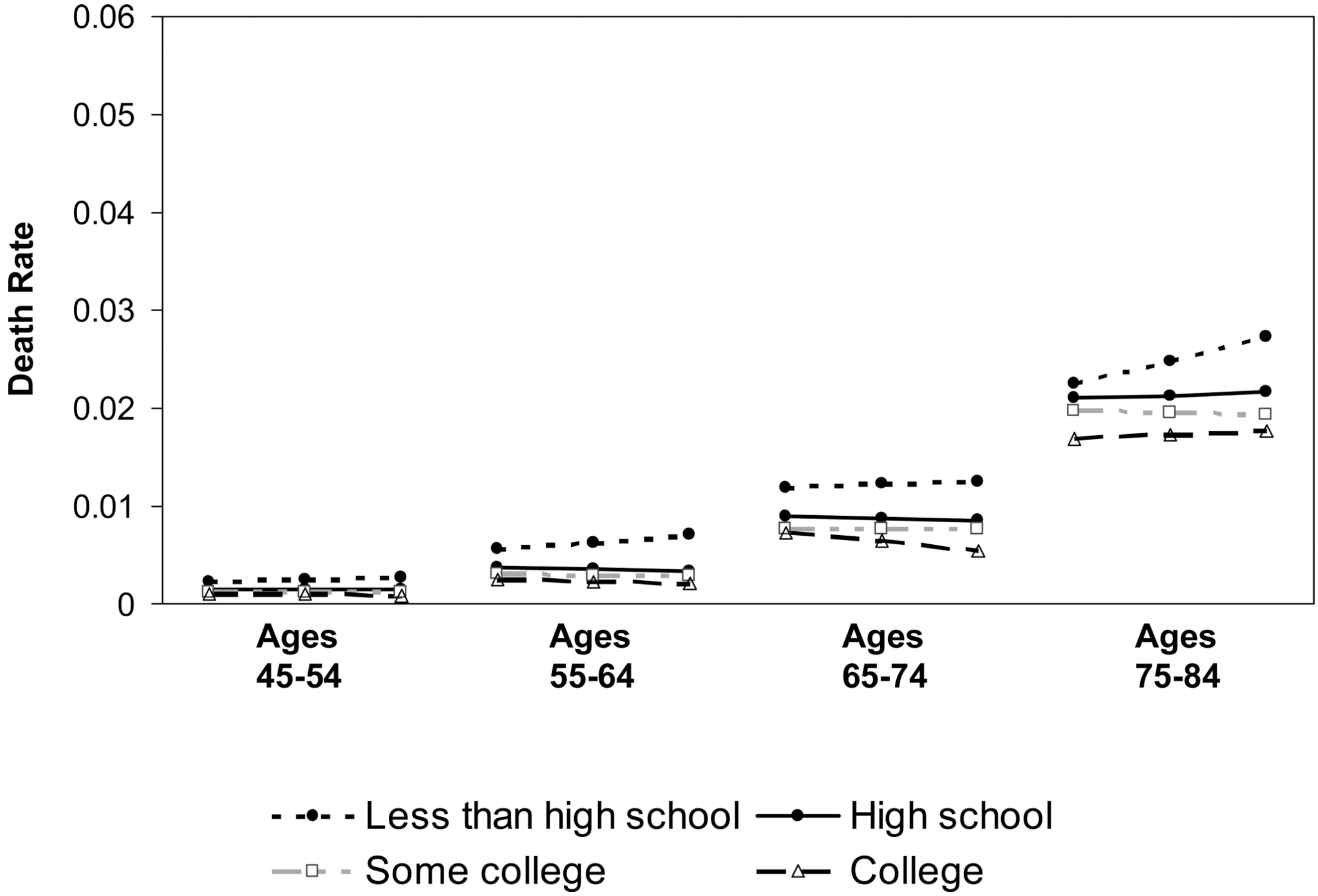

Figure 2.

Trends in Annual Death Rate for Non-Hispanic White Women across Three Time Periods (1986–1992, 1993–1999, 2000–2006) by Educational Attainment and Age Group

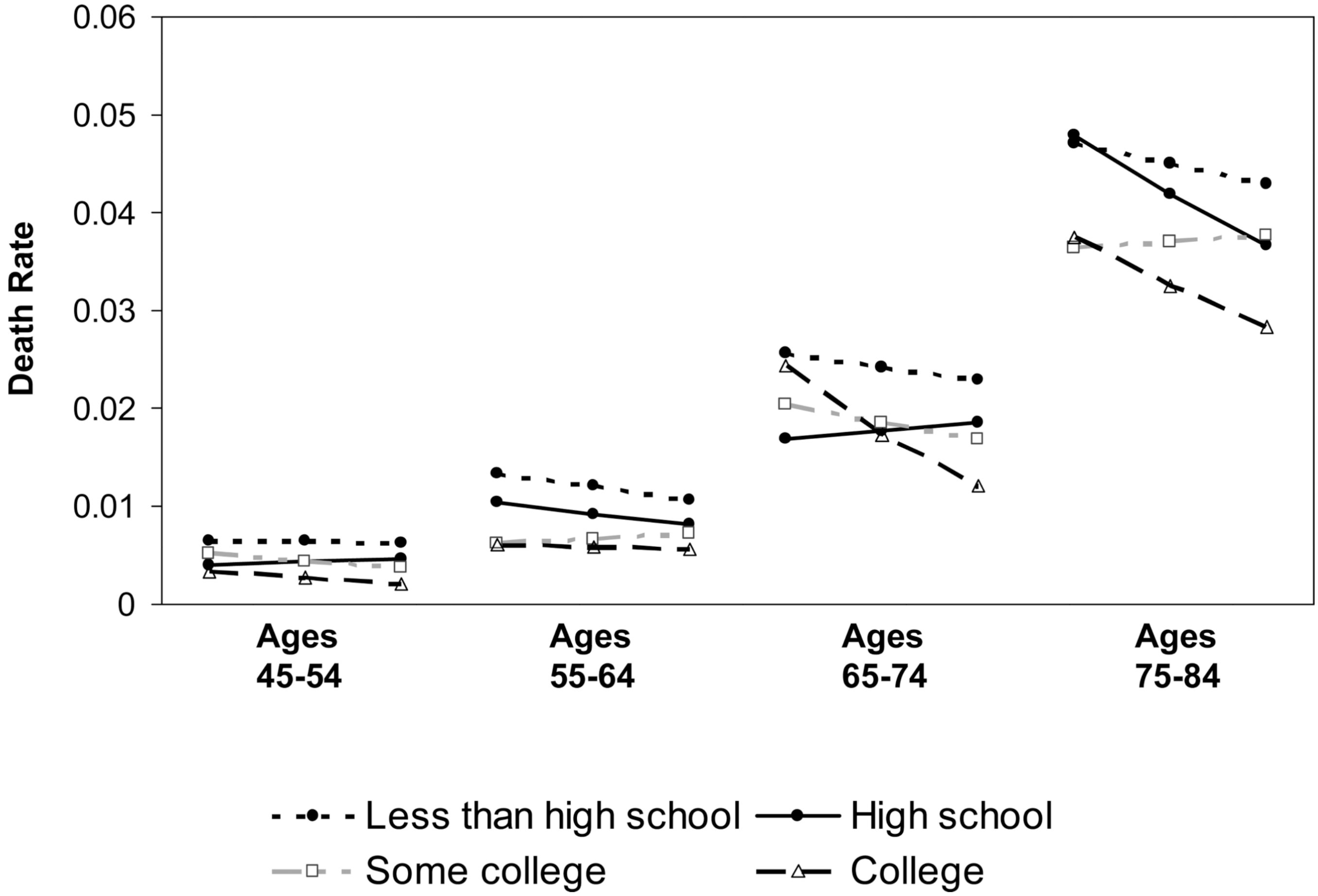

Figure 3.
Trends in Annual Death Rate for Non-Hispanic Black Men across Three Time Periods (1986–1992, 1993–1999, 2000–2006) by Educational Attainment and Age Group

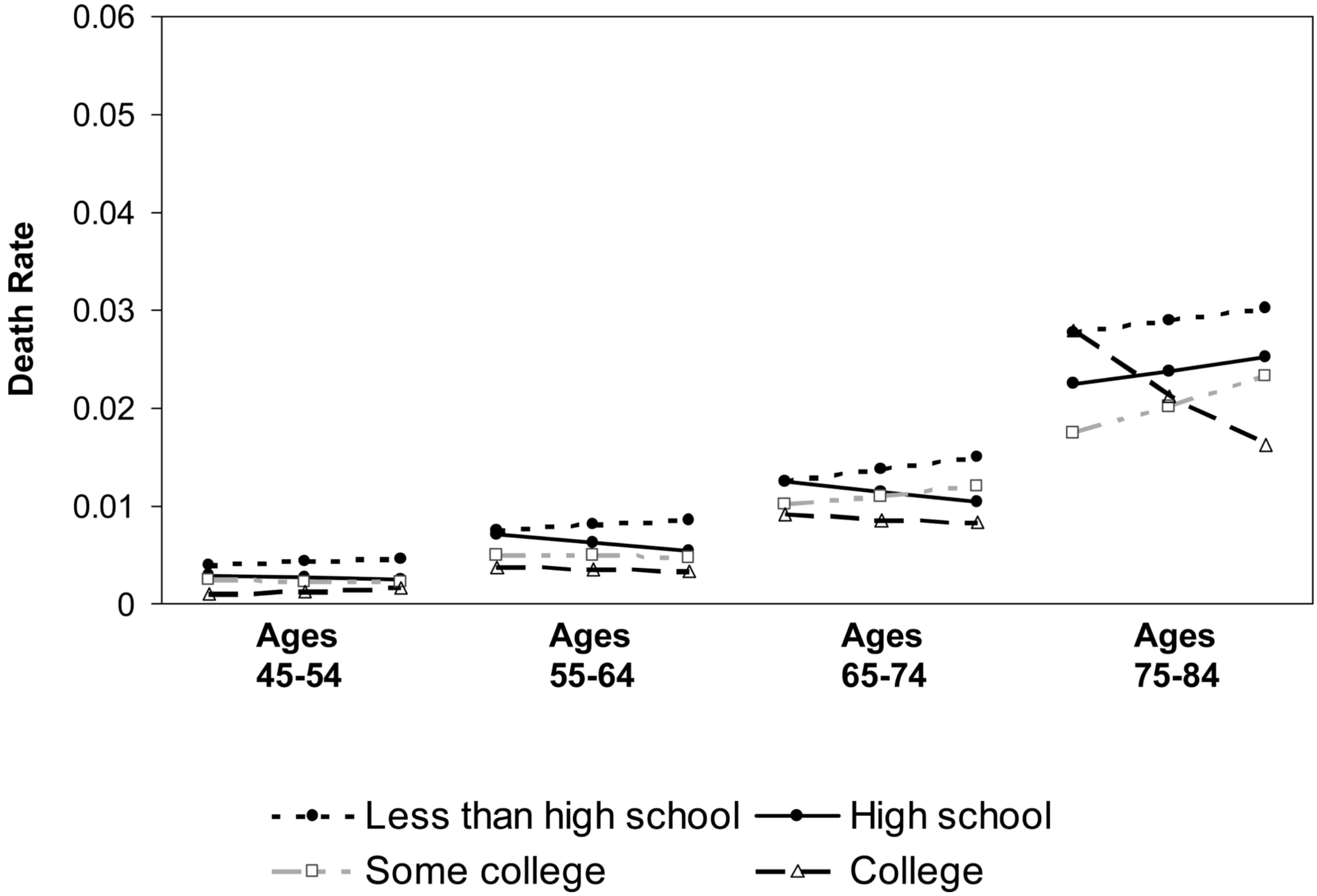

Figure 4.

Trends in Annual Death Rate for Non-Hispanic Black Women across Three Time Periods (1986–1992, 1993–1999, 2000–2006) by Educational Attainment and Age Group

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Table 1

Distribution of Educational Attainment[a] by Gender and Age among Non-Hispanic Whites Aged 45–84 within Three Time Periods

	Men			Women		
Education	**1986–1992**	**1993–1999**	**2000–2006**	**1986–1992**	**1993–1999**	**2000–2006**
Less than High School	*.267*	*.194*	*.140*	*.269*	*.198*	*.141*
45–54 age group	.155	.103	.088	.152	.099	.077
55–64 age group	.260	.180	.119	.236	.175	.116
65–74 age group	.342	.280	.200	.324	.257	.194
75–84 age group	.471	.357	.272	.463	.362	.267
High School	*.349*	*.342*	*.334*	*.441*	*.437*	*.407*
45–54 age group	.369	.325	.336	.464	.409	.365
55–64 age group	.359	.371	.322	.485	.473	.405
65–74 age group	.345	.344	.351	.443	.468	.462
75–84 age group	.270	.328	.330	.320	.400	.443
Some College	*.153*	*.185*	*.220*	*.157*	*.189*	*.234*
45–54 age group	.185	.227	.254	.192	.235	.276
55–64 age group	.146	.175	.225	.156	.184	.239
65–74 age group	.134	.146	.176	.138	.158	.193
75–84 age group	.115	.140	.165	.123	.141	.176
College or Higher	*.231*	*.280*	*.305*	*.133*	*.176*	*.219*
45–54 age group	.290	.344	.322	.192	.257	.282
55–64 age group	.237	.274	.334	.124	.167	.240
65–74 age group	.179	.230	.273	.094	.117	.152
75–84 age group	.144	.176	.233	.093	.097	.114
Number of Deaths	6,124	18,485	26,094	4,679	15,298	22,528
Number of Person-years	284,422	935,505	1,530,271	337,394	1,085,793	1,732,477

[a]Distributions reflect the percentage of person-years and are adjusted by the sample weights.

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Table 2

Distribution of Educational Attainment[a] by Gender and Age among Non-Hispanic Blacks Aged 45–84 within Three Time Periods

	Men			Women		
Education	**1986–1992**	**1993–1999**	**2000–2006**	**1986–1992**	**1993–1999**	**2000–2006**
Less than High School	*.520*	*.392*	*.286*	*.495*	*.378*	*.281*
45–54 age group	.352	.235	.180	.322	.219	.172
55–64 age group	.536	.405	.276	.489	.376	.260
65–74 age group	.672	.572	.441	.641	.533	.420
75–84 age group	.784	.685	.589	.759	.654	.549
High School	*.280*	*.329*	*.355*	*.311*	*.353*	*.362*
45–54 age group	.365	.389	.393	.407	.413	.386
55–64 age group	.272	.334	.360	.322	.370	.379
65–74 age group	.209	.257	.300	.232	.292	.334
75–84 age group	.123	.197	.240	.140	.220	.273
Some College	*.113*	*.154*	*.210*	*.106*	*.152*	*.215*
45–54 age group	.160	.207	.257	.152	.214	.274
55–64 age group	.112	.149	.204	.102	.144	.214
65–74 age group	.063	.093	.149	.066	.092	.143
75–84 age group	.046	.060	.095	.052	.066	.096
College or Higher	*.088*	*.125*	*.149*	*.089*	*.116*	*.141*
45–54 age group	.122	.170	.171	.119	.154	.167
55–64 age group	.080	.112	.159	.088	.110	.147
65–74 age group	.056	.077	.110	.062	.082	.104
75–84 age group	.046	.058	.077	.050	.060	.082
Number of Deaths	1,147	3,246	4,714	984	3,249	5,026
Number of Person-years	39,606	124,856	210,695	58,428	187,845	315,215

[a] Distributions reflect the percentage of person-years and are adjusted by the sample weights.

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Table 3

Poisson Regression Coefficients Predicting the ln(annual death rate) from Age, Time, and Education among Non-Hispanic White Men and Women Aged 45–84 in the United States, 1986–2006

	Men						Women					
	M1 45–84	M2 45–84	Ages 45–54	Ages 55–64	Ages 65–74	Ages 75–84	W1 45–84	W2 45–84	Ages 45–54	Ages 55–64	Ages 65–74	Ages 75–84
Intercept	−10.691**	−10.560**	−11.550**	−10.974**	−10.252**	−10.592**	−11.359**	−11.109**	−11.255**	−11.122**	−10.648**	−11.951**
Trend[a]	−0.070**	−0.122**	−0.022	−0.093*	−0.192**	−0.127**	0.019**	−0.078**	−0.179*	−0.124*	−0.144**	0.025
Age in Years	0.089**	0.089**	0.099**	0.092**	0.087**	0.091**	0.089**	0.089**	0.093**	0.088**	0.084**	0.098**
Education (College)[b]												
LTHS	0.677**	0.503**	0.880**	0.855**	0.398**	0.259**	0.648**	0.269**	0.455†	0.558**	0.315*	0.221*
HS	0.426**	0.297**	0.685**	0.442**	0.152†	0.184†	0.354**	0.161*	−0.009	0.290†	0.095	0.231*
SC	0.336**	0.189**	0.346	0.596**	0.175	−0.121	0.234**	0.052	−0.128	0.082	−0.091	0.202†
Trend × Education												
Trend × LTHS		0.071**	0.108	0.055	0.125**	0.076*		0.153**	0.261**	0.232**	0.167**	0.071†
Trend × HS		0.051*	0.015	0.051	0.104**	0.034		0.075**	0.219*	0.077	0.111*	−0.010
Trend × SC		0.058*	0.071	−0.047	0.078†	0.117**		0.072*	0.184†	0.102	0.144*	−0.035
Deviance												
Main effects model		405,077	54,798	82,162	125,432	142,350		359,312	37,762	61,850	103,387	155,924
Interaction model		405,065	54,795	82,156	125,419	142,339		359,271	37,754	61,834	103,376	155,908
χ^2		12**	3	6	13**	11*		41**	8*	16**	11*	16**
Deaths	50,703	50,703	4,638	8,275	15,686	22,104	42,505	42,505	3,027	5,807	11,743	21,928
Person-years	2,750,198	2,750,198	1,079,322	773,347	561,847	335,682	3,155,665	3,155,665	1,142,098	832,437	674,795	506,335

Notes: Reference groups in parentheses. The χ^2 statistic is the difference in the deviances for the main effects and interaction models. A statistically significant value indicates that the interaction model provides a better fit to the data.

[a]Trend is a 3-level ordinal variable. Trend=1 for 1986–1992, trend=2 for 1993–1999, and trend=3 for 2000–2006.

[b]LTHS = less than high school; HS = high school; SC = some college.

†p<0.10;

* p<0.05;

** p<0.01 (two-tailed tests for Poisson coefficients; one-tailed tests for model χ^2)

Table 4

Poisson Regression Coefficients Predicting the ln(annual death rate) from Age, Time, and Education among Non-Hispanic Black Men and Women Aged 45–84 in the United States, 1986–2006

	Men						Women					
	M1 45–84	**M2 45–84**	**Ages 45–54**	**Ages 55–64**	**Ages 65–74**	**Ages 75–84**	**W1 45–84**	**W2 45–84**	**Ages 45–54**	**Ages 55–64**	**Ages 65–74**	**Ages 75–84**
Intercept	−9.229**	−8.878**	−10.303**	−8.810**	−7.561**	−9.543**	−9.812**	−9.560**	−11.312**	−10.453**	−9.415**	−8.672**
Trend[a]	−0.060**	−0.196**	−0.215	−0.026	−0.351**	−0.141	0.024	−0.074	0.220	−0.056	−0.048	−0.269*
Age in Years	0.071**	0.071**	0.096**	0.062**	0.060**	0.080**	0.070**	0.070**	0.085**	0.082**	0.068**	0.067**
Education (College)[b]												
LTHS	0.611**	0.262	0.489	0.890*	−0.251	0.129	0.681**	0.330	1.446**	0.591	0.179	−0.320
HS	0.385**	−0.043	−0.064	0.638	−0.763**	0.237	0.402**	0.348	1.235*	0.696†	0.375	−0.531
SC	0.285**	−0.090	0.412	−0.069	−0.439	−0.185	0.300**	0.008	1.065†	0.224	−0.005	−0.872
Trend × Education												
Trend × LTHS		0.136†	0.199	−0.086	0.296**	0.097		0.138†	−0.139	0.116	0.140	0.314*
Trend × HS		0.167*	0.275	−0.093	0.396**	0.007		0.020	−0.264	−0.069	−0.046	0.322*
Trend × SC		0.146†	0.049	0.101	0.258†	0.157		0.113	−0.257	0.052	0.128	0.410*
Deviance												
Main effects model		54,755	12,076	13,166	16,175	13,296		54,652	10,288	12,521	15,532	16,269
Interaction model		54,751	12,072	13,164	16,167	13,294		54,643	10,285	12,517	15,526	16,265
χ^2		4	4	2	8*	2		9*	3	4	6	4
Deaths	9,107	9,107	1,349	1,880	2,889	2,989	9,259	9,259	1,265	1,837	2,722	3,435
Person-years	375,157	375,157	157,225	106,635	73,274	38,023	561,489	561,489	226,452	154,041	112,126	68,870

Notes: Reference groups in parentheses. The χ^2 statistic is the difference in the deviances for the main effects and interaction models. A statistically significant value indicates that the interaction model provides a better fit to the data.

[a] Trend is a 3-level ordinal variable. Trend=1 for 1986–1992, trend=2 for 1993–1999, and trend=3 for 2000–2006.

[b] LTHS = less than high school; HS = high school; SC = some college.

† p<0.10;

* $p<0.05$;

** $p<0.01$ (two-tailed tests for Poisson coefficients; one-tailed tests for model χ^2)