Published in final edited form as:

Popul Res Policy Rev. 2013 December 1; 32(6): . doi:10.1007/s11113-013-9299-0.

The Nonlinear Relationship between Education and Mortality: An Examination of Cohort, Race/Ethnic, and Gender Differences

Bethany G. Everett,
University of Illinois-Chicago

David H. Rehkopf, and
David, H. Rehkopf, Assistant Professor, Medicine, Stanford University, 300 Pasteur Drive, Room 2102, MC 5110, Stanford, CA, 94305

Richard G. Rogers
Richard G. Rogers, Professor, Sociology and Institute of Behavioral Science, 483 UCB, Boulder, CO 80309-0483

David H. Rehkopf: Drehkopf@stanford.edu; Richard G. Rogers: Richard.Rogers@colorado.edu

Abstract

Researchers investigating the relationship between education and mortality in industrialized countries have consistently shown that higher levels of education are associated with decreased mortality risk. The shape of the education-mortality relationship and how it varies by demographic group have been examined less frequently. Using the U.S. National Health Interview Survey-Linked Mortality Files, which link the 1986 through 2004 NHIS to the National Death Index through 2006, we examine the shape of the education-mortality curve by cohort, race/ethnicity, and gender. Whereas traditional regression models assume a constrained functional form for the dependence of education and mortality, in most cases semiparametric models allow us to more accurately describe how the association varies by cohort, both between and within race/ethnic and gender subpopulations. Notably, we find significant changes over time in both the shape and the magnitude of the education-mortality gradient across cohorts of women and white men, but little change among younger cohorts of black men. Such insights into demographic patterns in education and mortality can ultimately help increase life expectancies.

Keywords

Education; Mortality; Race/ethnicity; Gender; Semiparametric modeling; NHIS

Background

A longstanding body of work has documented the graded relationship between socioeconomic status (SES) and mortality in the United States (Adler and Rehkopf 2008; Antonovsky 1967; Hummer and Lariscy 2011; Kitagawa and Hauser 1973; Preston and Elo 1995; Rogers et al. 2010). SES has been described as the fundamental cause of health disparities in the United States because income, power, and knowledge allow people to adapt to new health-related information to improve their health (Link and Phelan 1995; Phelan et al. 2004). Education is a key indicator of SES and has certain advantages over

income as a marker of SES, as it is established early in life and is not subject to the same fluctuations over the lifecourse, particularly in old age (Hummer and Lariscy 2011).

Four decades of research examining the relationship between education and mortality in the U.S. have demonstrated that as education increases, mortality risk decreases and life expectancy increases (Hummer and Larsicy 2011; Kitagawa and Hauser 1973; Preston and Elo 1996; Rogers, Hummer, and Nam 2000). But continued investigation of this relationship is warranted, as new research suggests that life expectancies among persons with less than 12 years of education may actually be decreasing in the United States (Olshansky et al. 2012). Moreover, it is less clear how the relationship varies by race/ethnicity, gender, and age cohort, although recent studies suggest that these factors affect the health returns to education (Jemal et al. 2008; Meara, Richards, and Cutler 2008; Montez et al. 2009; Olshansky et al. 2012; Zajacova 2006; Zajacova and Hummer 2009). Existing research has used several types of models, including examining education as a linear term or a categorical term, and other novel approaches that allow the education-mortality slope to vary across levels of educational achievement. Using the 2006 National Health Interview Survey-Linked Mortality Files (NHIS-LMF), this study builds upon the existing research by using spline-based semiparametric Cox proportional hazard models. The primary advantage of this modeling approach is that we do not have to rely on a prespecified functional form of the relationship between education and mortality as have previous studies. Identifying potential nonlinear relationships is important for understanding the etiology of the association and for focusing resources to reduce education-related disparities in mortality.

Cohort Differences in the Relationship between Education and Mortality

The economic returns to education historically have not been as high as they are today. The “high school movement” of the 1940s greatly increased the number of available skilled workers, and thus competition in the labor market. This increase in supply, in the absence of a uniform increase in demand, decreased wage inequality between those with high school degrees and those without (Goldin and Katz 1991, 2011). Since the 1970s, changes in technology and the labor force have increased the amount of time that one must spend in school to get a white-collar job. But there has been no increase in the supply of college-educated persons. A recent review of trends in educational achievement by race/ethnicity shows that average levels of education have stalled if not decreased among males born between 1965 and 1982 (Everett et al. 2011). This stagnation in supply coupled with growing demand for college-educated workers has increased the returns to education for those in the most highly educated groups in younger cohorts (Card and Lemieux 2011). Fundamental cause theory (Link and Phelan 1995) suggests that the recent growth in socioeconomic disparities should also be reflected in population-level trends in health, as those with more education can continue to improve their health and take advantage of new health-related technologies, while those with less education have less access to health-related amenities, such as gyms and health clubs, as well as health insurance and treatment.

Several studies have shown that the education-mortality gradient has become steeper in recent decades (Masters, Hummer, and Powers 2012; Pappas et al. 1993; Rogers et al. 2010; Rogers, Hummer, and Everett 2012; Steenland, Henley, and Thun 2002; Zajacova 2006; Zajacova and Hummer 2009). Net of age effects, the health gap between the highly and the less highly educated is larger among younger cohorts (Lynch 2003), largely because of a growing disparity in preventable causes of death (Masters et al. 2012; Miech et al. 2011).

Race/Ethnic and Gender Differences in the Relationship between Education and Mortality

An important supposition of fundamental cause theory is that persons with higher levels of education will be better able to get and respond to new health-related information and

implement the appropriate behaviors and treatments to improve their health and lengthen their lives (Link and Phelan 1995; Phelan et al. 2004). Thus, it is important to examine how the relationship of SES to health varies for different populations and has changed over time (Kunitz 2007).

While black labor force participation varied over the twentieth century, at times reaching that of whites (Bonacich 1976), or, in the case of black women, exceeding that of whites (Goldin 1977), substantial structural barriers to acquiring education and finding skilled employment exist for women and black Americans born in the first half of the twentieth century. The civil rights movement was important for improving the ability of black Americans to increase both the quantity and the quality of their education. Unfortunately, the boom of educational progress that occurred shortly after the civil rights act has slowed if not stalled (Donohue and Heckman 1991; Neal 2006; Everett et al. 2011). Hanushek and Rivkin (2006) have shown that between-school gaps in achievement can be in large part explained by differences in indicators of school quality, such as teacher quality, which tend to coincide with racial gaps. Discrimination, net of educational attainment, may result in lower incomes and increased difficulty finding employment for non-Hispanic black Americans (Arrow 1973; Altonji and Blank 1999; Bertrand and Mullainathan 2004; Phelps 1972). Further, there is some evidence that higher levels of stress and discrimination are important predictors of cardiovascular health that may undermine the benefits of education for race/ethnic minorities in the United States (Brunner 1997; Krieger et al. 1993).

Studies that have examined black-white differences in the relationship between education and mortality have found a more strongly graded relationship among whites. Jemal et al. (2008) showed that the mortality gap between the lowest and highest levels of education was fourfold among white males and just threefold among black males. Meara et al. (2008) showed that life expectancy has increased steadily among white men, but has lagged and even stalled in recent years among more educated black males.

Changes in the returns to education have also occurred for women. Women's college attendance rates hit an early peak in the 1940s, but dropped dramatically in the 1950s and 1960s (Jacobs 1996). During the period of the civil rights movement dramatic changes in labor force participation gender norms encouraged women's educational achievement (Cancian and Gordon 1988). As a result, in recent years, women's rates of educational achievement have increased substantially, so that among recent cohorts, women now have higher average education than men (Buchmann, DiPrete, and McDaniel 2008; Everett et al. 2011).

These changes should be reflected in the relationship between education and mortality. Indeed, several recent studies have found that across cohorts, women's education-mortality gradient is increasingly steeper compared to men's (Zajacova 2006). Zajacova and Hummer (2009) found that the education-mortality gradient was steeper for men than for women among non-Hispanic whites born between 1906 and 1945, but not among whites born between 1946 and 1965, and found no differences between black men and women for any cohort. Thus, we expect that the education- mortality gradient will be steeper among younger compared to older cohorts, and the sharpening over time will be most dramatic among non-Hispanic whites and women.

Nonlinearity in the Relationship between Education and Mortality

While it is common to assume a nonlinear relationship between health outcomes and income, few studies have systematically examined the shape of the education-mortality slope. Rather, the relationship is most often estimated by categorizing years of schooling into a series of dummy variables (see Kitagawa and Hauser 1973; Molla, Madans, and

Wagener 2004; Rogers et al. 2012) or using a linear predictor (Elo, Martikainen, and Smith 2006; Lynch 2006; Preston and Elo 1995; Zajacova 2006). Using a linear term is problematic because much research has shown that the relationship between years of education and mortality risk is nonmonotonic (Rosse and Mirosky 2003; Backlund, Sorlie, and Johnson 1999; Montez et al. 2012). While using dummy variables for groups of more than one year of education skirts the issue of nonlinearity, it often obscures potential differences in the shape of the mortality curve as well as potential threshold differences between populations. Mirowsky and Ross (2003) have suggested that the slope is linear with a “step-change” at 12 years of education, but not for college degrees.

More recently Montez and colleagues (2012) used 13 different logistic regression model specifications to examine the functional form of the education-mortality gradient. They found that there was a linear decline in the slope between 0 and 11 years of education, and a “step-change” reduction at specific years that mark degree completion followed by a steeper linear slope. Neither of these studies found a step-change at 16 years of education. While theirs is the most thorough examination to date, their models are not semiparametric and thus are constrained by the specified model. Other studies that have used semiparametric nonlinear models to explore the relationship among income, biomarkers of coronary heart disease, and overall mortality (Rehkopf et al. 2008, 2010) have demonstrated that this method improves model fit and better describes the shape of the relationship between SES and health outcomes. Semiparametric models also allow a formal statistical test of nonlinearity and are generally more stable with smaller sample sizes than models with dummy variables and slopes; thus they allow examination of relationships within population subgroups (Ramsey and Ripley 2007; Wood 2006).

Aims

The analyses presented in this paper improve upon existing research on the relationship between education and mortality by using updated mortality information; disaggregating our analysis by cohort, race/ethnicity, and gender; and employing a more flexible modeling approach. The use of a penalized spline on education allows us to use all available information on years of education and create more stable estimates for each year of education. We test two research questions: are there substantially nonlinear relationships between education and mortality that can be more usefully described by a semiparametric modeling approach, and does the education and mortality relationship vary among cohorts by race and gender?

Methods

Data and Measures

We employ data from the public-use National Health Interview Survey-Linked Mortality Files (NHIS-LMF). NHIS-LMF links the 1986 through 2004 NHIS to mortality using the National Death Index through the year 2006 (NCHS 2005, 2010). This data set is ideal for our purposes because it is extremely large, is nationally representative, and allows us to examine the link between education and mortality among several subpopulations. To provide a public-use version of these data, NCHS perturbed the dates or cause of death for a small number of records to ensure that individuals could not be identified. Lochner and colleagues (2008) demonstrated that the public-use and restricted data sets produce equivalent results for overall mortality. We restrict our sample to respondents ages 35 to 65 years old, approximately the working age population, for three reasons: (1) persons over 35 are likely to have completed their education, (2) these restrictions result in overlaps in the ages of respondents across cohorts, and (3) this age range limits the biases that may be introduced by differences in the causes of death between young and old populations. We

also restrict our analysis to respondents who identify as non-Hispanic white (hereafter “white”) or non-Hispanic black (hereafter “black”), in order to have enough respondents to examine gender and cohort differences within race/ethnic groups. Focusing on non-Hispanic whites and blacks results in a sample size of 539,770 respondents, of whom 56,432 died during the follow-up period.

We use Cox proportional hazard models to examine mortality by race and gender. Our dependent variable is observed to the event “death,” with right censoring for people who survived the follow-up period. Education is coded as a continuous variable that ranges from 0 to 18 or more years of education (that is, it is top-coded at 18 or more years). NHIS changed the coding strategy for education between 1996 and 1997 to make educational attainment a categorical variable, asking whether respondents had completed 1-12 years of school, graduated from high school, or obtained an associate's, bachelor's, master's, professional, or doctoral degree. (GEDs are coded as 12 years of education, a limitation of our study given the body of work that suggests differences in the health returns to GEDs compared to high school diplomas [Rogers et al. 2010].) To adjust for this coding change, we converted the categories into comparable years of education and include a dummy variable for whether persons were surveyed before 1997 or in or after 1997. We graphed the mean education level by year and did not discern any breaks between 1996 and 1997 due to coding changes, which suggests that the measurement change has little effect on the study results. Other studies have used a similar recoding strategy (Everett et al. 2011; Zajacova and Hummer 2009). Additionally, we include a control for the year of the survey, which ranges from 1986 to 2004.

Method

We use Cox proportional hazard models with a penalized spline on the education term. The penalized spline model has substantial advantages over simpler models where the user specifies one or more inflection points or one or more categories to model the relationship. Categories must be prespecified by the user, and they limit power to detect differences. The primary emphasis of this analysis is to determine the extent of non-linearity in different population groups. The penalized spline algorithm allows our models to adjust for potential nonlinearity by fitting a model in two steps. First, the relationship between education and mortality is estimated using a large number of knots to form splines, or separate slopes for the relationship between each year of education and mortality risk. That is, at each year of education, the slope is allowed to vary. Second, the models then uses Akaike Information Criterion (AIC) fit to smooth the spline using a penalized smoothing parameter that balances over-fitting the model with excessive degrees of freedom with a more restricted model that still allows for nonlinear variations in the relationship between education and mortality. For example, if the relationship between education and mortality were a straight line, allowing for changes in slopes at every year of education would unnecessarily increase the degrees of freedom used in the model and result in an overfitting of the data. Ruppert, Wand, and Carroll (2003) have shown that this method is insensitive to the number of knots initially chosen. The penalized spline approach, therefore, allows for the results to reflect nonlinear variations in the relationship between education and mortality while optimizing model fit (Eilers and Marx 1996; Hurvich, Simonoff, and Tsai 1998).

We estimate a series of Cox proportional hazard models with different specifications to ensure that we are correctly modeling the relationship between education and mortality, and to examine how robust our results are to different model specifications. Thus, we include results from Cox proportional hazard models with the following model specifications: education as a linear term, education as a categorical term (less than 12 years of education, 12 years of education, and greater than 12 years of education), education with a penalized spline on the education term, education with a penalized spline on the education term with a

Generalized Estimating Equation (GEE) specification, education with a penalized spline with a frailty model gamma distribution, and education with a penalized spline with a frailty model with an inverse-Gaussian distribution.¹ GEE and frailty models allow for individual-level heterogeneity from the population-level estimate of the hazard slope. The frailty model can therefore be thought of as a random effect that allows for individuals to deviate from the mean hazard to either raise or lower their individual risk from the population. The GEE specification does not specify a functional form of the random effect. Functional forms of the random effect are specified in frailty models: the gamma distribution assumes that the individual level effect disappears over time, whereas the inverse-Gaussian effect assumes that the individual level effect stays constant over time.

Following Kom, Graubard, and Midthune (1997), we use age to indicate the time to death, which ensures that our mortality analyses are age-adjusted. We adjust for sample weights and strata to account for NHIS's complex sampling frame. All analyses are completed in R version 2.15.1 (R-Development-Core-Team). The proportional hazard models use the `coxph()` commands, and penalized splines are fit with the `pspline()` (Ramsey and Ripley 2007). Thus, we estimate the hazard for the i^{th} individual as

$$\lambda_i(t) = \lambda_0(t) \exp(Z_i(t)\beta_z + s(X_i(t))),$$

where s is the smoothed function containing multiple slopes and knots as specified by the degrees of freedom of the smooth for our primary exposure of interest (education).

For each of these models, we present the education hazard ratio, the AIC fit statistic, and the ROC area statistic, and indicate whether the model satisfied the Cox proportionality assumptions. The AIC statistic provides an estimate of the overall model fit. The ROC (receiver operator characteristic) area statistic provides an estimate of the predictive power of our models, using the R statistical package “Verification” (NCAR 2012), by giving an estimate of how well our model correctly classifies respondents as deceased. The ROC curve plots the sensitivity (the number of deaths that were correctly predicted as deaths) over 1-specificity (the number of people predicted as being dead who did not in fact die during the follow-up period). Data points are the ratio of the two estimates. The area under the curve (AUC) statistic therefore provides an estimate of predictive accuracy that ranges between 0 and 1, where 1 corresponds to 100% accuracy, and .5 corresponds to roughly a 50/50 chance of accurately predicting mortality; the more area under the curve, the more predictive power the model has. Models that have ROC statistics between .5 and .6 are considered to have failed the predictive test. We also include information on whether the models satisfy the proportional hazard assumptions using the Schoenfeld residuals test. Taken together, this information allows us to assess model fit across several specifications.

We stratify our analysis by race/ethnicity, gender, and the following birth cohorts: 1921–1936, 1937–1953, and 1954–1969. The Cox proportional hazard models using a nonlinear spline on the education term are additionally presented in graphical form, and the predicted values for years of education are presented in table form. To quantify whether the education-mortality curves differ by cohort, race, or gender, we specify several models that use interaction terms to examine whether education curves differ by race/ethnicity or gender within cohorts; whether black and white education curves differ by gender within cohorts;

¹We also estimated Cox proportional hazard models with penalized spline on the education term with both accelerated failure Weibull distributions and accelerated failure logged distributions. These two specifications, however, yielded very poor model fit and are therefore excluded from the results presented in the paper.

and whether education curves by race and gender change across cohorts. For example, to assess whether the education-mortality curve of white males differs significantly between cohorts 1 and 2, we restrict our analysis to white males born between 1921 and 1953 and regress mortality on a nonlinear education term, a dummy for the cohort 1937–1953, and a nonlinear interaction term between education and cohort 1937–1953. We specified a priori that we would define population differences as those where the interaction terms were significant at $p = 0.10$. Table 3 presents the hazard ratios and p -values for the interaction terms.

Results

Descriptive Statistics

Table 1 provides descriptive information on the distribution of education and the percent dead in education categories by race/ethnicity, cohort, and gender. Among the total population, for both race/ethnic groups and both genders, there is a steady decrease across cohorts in the proportion of respondents who report 8 or fewer years of education. This decrease is most dramatic among black men, where 30.1% black men born between 1921 and 1936 report 8 or fewer years of education compared to just 2.5% born between 1954 and 1969. However, the results also suggest stark differences in educational achievement by race/ethnicity, even among the youngest cohort. This is perhaps most striking at the higher levels of achievement: 30% of white men and 23% of white women have 16 years or more of education, compared to 15% of black men and 16% of black women.

For all groups, the results show larger proportions of deaths among the less educated for all cohorts. But the difference in percent dead between the least and most educated is largest among the youngest cohort. For example, the difference in percent dead between white males with zero to 8 years of education and those with 19 to 20 years of education is 54% for males born between 1921 and 1936, but 82% for males born between 1954 and 1969. For white females, the percent difference jumps from 58% to 83%, for black males from 48% to 77%, and for black females from 50% to 88%.

Nonlinear Effects of Education on Mortality by Cohort

Table 2 presents the AIC statistic, ROC area statistic, and whether the model passes the Schoenfeld residuals test for proportionality of the education term for the Cox proportional hazard models. Smaller values of the AIC indicate better model fit, taking into account that the nonparametric models have a greater degree of freedom. This metric allows us to evenly compare models of different types, and allows a simpler model (fewer degrees of freedom) with a slightly worse fit to the data to have a lower (better) AIC. In models with large AIC values, even AIC reductions as small as 10 indicate important improvement in fit (Burnham and Anderson 2004). Larger values for the ROC statistic indicate better predictive power. This table shows that with few exceptions, the basic nonlinear model is the most appropriate model for analyzing the relationship between education and mortality. While the Cox models with frailty and Gaussian distribution often have better predictive power, the AICs are much larger compared to the other model specifications and have higher rates of failure of the proportionality assumptions. For black men born between 1921 and 1936, the categorical model is a better model fit than the nonlinear model, and for black men born between 1937 and 1953 and black women born between 1954 and 1969, the categorical and nonlinear models both satisfy the proportionality assumption, with one model providing a better AIC and the other model providing a better ROC statistic. Thus, our results suggest that the best-fitting alternative model is the categorical model.

Having demonstrated the advantages of the nonlinear model for most groups analyzed, we present the results from these models in graphical form. The dotted horizontal line at the hazard ratio of one represents the mean mortality risk for the population examined; points on the curve above one represent an elevated mortality risk and points below represent a decreased mortality risk. The solid line represents the mortality risk as a hazard ratio, and the dashed lines on both sides represent the 95% confidence interval around the mortality risk.

Males—There is a slow linear decline in the education-mortality relationship between zero and 9 years, after which there is a steady decline in the mortality risk that crosses the mean risk at 13 years of education. Table 3 shows that across cohorts, the hazard ratio associated with low levels of education increases every year, although among white men, there is little change across cohort in the education benefit for high levels of education. That is, among younger generations those with little education are increasingly at relatively higher mortality risk, a finding that is in line with other work showing a growing education-mortality disparity. The interactions in Table 4 show that the education-mortality slope becomes steeper across cohorts; this finding is represented in Figure 1 as well.

Among black males born between 1921 and 1936, the relationship between education and mortality is almost flat until 12 years of education, after which there is a slight decrease in mortality risk, so that black men with 18 years of education are 63% less likely to die in the follow-up period compared to the mean risk for this group. A significantly steeper education-mortality gradient emerges among black males born between 1937 and 1953, crossing the mean risk at 14 years of education. The mortality risk curve for black men born between 1954 and 1969 suggests an additional increase in mortality risk at the lowest levels of education (HR=2.35); however, there are wide confidence intervals around the low levels of education among this cohort of black men. Moreover, the results from interactions suggest that the slope for black men born between 1954 and 1969 is not significantly different from the previous cohort. Within all cohorts, the education-mortality gradient among white males is steeper than the gradient among black males. The results also suggest that while educational disparities in mortality are emerging among black males, their emergence is both delayed and less pronounced compared to the change among whites. For example, among white men with zero years of education there is a 67% increase in the hazard ratio from the oldest to the youngest cohort, but among black men there is only a 46% increase in risk.

Females—The education-mortality curve for women born between 1921 and 1936 is almost flat for women with zero to 7 years of education and then begins to decrease more rapidly from 8 to 14 years of education before leveling off again. The mortality curve crosses the mean risk for this cohort at 12 years of education. Women with 18 years of education are 67% less likely to die during the follow-up period compared to the mean risk, while women with zero years of education are 2.15 times more likely to die. As is true for white males, the education-mortality gradient of each subsequent cohort differs statistically from that of the previous cohort, so that women born between 1954 and 1969 with zero years of education are almost 5.8 times as likely to die over the follow-up period compared to the mean risk; however, the hazard ratio for women with high levels of education remains relatively stable across cohorts. The results for white women differ slightly from those of previous studies that showed two slopes with a step change at 12 years of education (Mirowsky and Ross 2003; Montez et al. 2012), as they suggest three slopes, with changes at roughly 9 and 14 years.

A similar steep mortality gradient emerges among black females, although its emergence is delayed until the cohort of women born between 1937 and 1953. The education-mortality

gradient that emerges among the 1954–1969 birth cohort is significantly steeper than that for the previous cohort, as is the 1937–1953 mortality curve compared to the 1921–1936 mortality curve. Indeed, the slope crosses the mean mortality risk at 12 years for the oldest cohort, 13 years in the subsequent cohort, and 14 years in the youngest cohort. In the youngest cohort, compared to the mean risk, the least educated black women are 3.7 times as likely to die during the follow-up period. Black women's education-mortality gradient is significantly flatter than that of white females in all cohorts; however, the hazard ratio for the lowest levels of education increases roughly 65% for both black and white women from the oldest to the youngest cohort. This suggests that while the slope is steeper at baseline for white women, the educational disparity in mortality is emerging at a similar rate for both white and black women.

Table 4 shows that the mortality risk curve is significantly steeper in each subsequent cohort, except for black males. The gradient is also steeper for white men than for black men, and for white women than for black women. Within race/ethnic groups, white males have a significantly steeper slope than do women in the oldest cohort only, and black women have a steeper slope than do men in the two oldest cohorts.

Because our results span several years and include many age groups, we examined whether the results held across causes of death, as the observed differences in the education-mortality relationship could simply be due to differences in age-related processes related to mortality. Following a coding scheme by Phelan et al. (2004), we examined whether the education-mortality gradient varied between causes of death with low and high preventability. The results, not shown but available upon request, demonstrated that the education-mortality gradient becomes steeper with each subsequent cohort for causes of death with both low and high preventability. This result holds for all demographic subgroups.

Discussion and Conclusion

Continuing to investigate the relationship between education and mortality remains important, particularly in light of the growing education-mortality divide (Masters et al. 2012; Olshansky et al. 2012; Rogers et al. 2012; Steenland et al. 2002; Zajacova 2006; Zajacova and Hummer 2009). Our results show that a nonlinear approach provides the best model fit for most groups analyzed and reveal education-mortality slopes that differ slightly from those found by previous studies. First, our results suggest that the slope change occurs at 9 years of education (rather than 12, as Montez et al. [2012] have found)—that is, at the transition from middle to high school. The education-mortality curves vary in shape for all groups analyzed and suggest an increasing educational disparity in mortality among younger cohorts, except for black men, where there is no significant change in the education-mortality slope in the two most recent cohorts. Across cohorts, the growing educational disparity in mortality appears to be driven primarily by increases in mortality risk at the lowest levels of education, rather than increasing mortality benefits at the highest levels of education.

Among white males, our results show that the education-mortality gradient emerges over time, with each cohort experiencing higher mortality penalties at the low ends of educational achievement, but relatively small reductions in mortality risk among those with high education. The results suggest that the slope between zero and eight years of education is the primary driver of the growing educational disparity in mortality, with nine years of education a critical point after which mortality risk begins to drop in a linear fashion through the education levels.

The changing relationship between education and mortality for white males with low levels of education across cohorts may be due to the changes in the returns to education. In the

early part of the twentieth century, the largely industrial economy provided a variety of high-wage employment opportunities that did not require high levels of post-high school education. Beginning in the 1970s, the U.S. labor market underwent radical changes characterized by increasing inequality and a segmented labor market with a large secondary low-skill, low-wage demand, and a smaller high-skill, high-wage specialization (Levy and Murnane 1992). Globalization, increasing vertical specialization, and a shift to a largely service-based economy have all decreased employment opportunities for people with little education (Fuchs 1968; Howell and Wolff 1991; OCED 2000). Further, the stagnation in the supply of highly educated people despite increased the gap in returns to education among younger cohorts (Card and Lemieux 2011). All of these factors have increased white men's mortality penalty for low levels of education.

Among black men, the delayed emergence of the education-mortality gradient and the lack of significant difference in its slope between men born between 1937 and 1953 and those born between 1954 and 1969 illustrates continued SES disadvantage. There were substantial barriers to acquiring education for black males for several decades, and persistent differences in school quality continue to systematically undermine black Americans (Hanushek and Rivkin 2006; Neal 2006). The economic outlook for black males, in fact, has actually worsened since the 1970s, with increasing disparities in unemployment and labor force participation (Holzer 2009), and lower levels of education than for black men born in the 1940s and 1950s (Everett et al. 2011). The increasing economic disparities have been in part attributed to high rates of incarceration among black males and, among those who have been released from prison, a reluctance of employers to hire people with criminal records (Holzer, Raphael, and Stoll 2004). Illegal activities, such as drug trade in the 1980s, both provided a viable economic alternative to education and decreased the likelihood of continued education for men caught and imprisoned (Fryer et al. 2005).

The extensive and continued history of racism in the United States may undermine black men's health returns to education in several important ways. First, high levels of both structural and interpersonal discrimination are substantial barriers to improving SES among black males (Arrow 1973; Altonji and Blank 1999; Bertrand and Mullainathan 2004; Phelps 1972). Discriminatory hiring practices may limit their access to income, insurance, and a variety of goods and services that may be used to improve health, even if they have achieved high levels of education. Moreover, discrimination may reduce the motivation to pursue high levels of education. Second, education is posited to improve health by affecting a variety of social psychological factors such as the individual's feelings of self-efficacy, problem solving, and feelings of personal control and mastery (Mirowsky and Ross 1998; Ross and Wu 1995). These educational benefits may be undermined for black men by continued exposure to discrimination and stigma, which are linked to poorer physical health outcomes (Kreiger 1990; Lewis et al. 2009; Ryan, Gee, and Laflamme 2006; Williams, Neighbor, and Jackson 2009).

Among women, our results show an emerging education-mortality gradient in the oldest cohort that becomes significantly steeper in each subsequent cohort. Women born between 1921 and 1936 faced substantial barriers to acquiring high levels of education, and even when they did, education did not necessarily translate into high-quality employment or high wages, and in turn better access to health care. Beginning in the 1960s, the political climate in the United States began to change, and several policies, including the Civil Rights Act in 1964 and the implementation of Title IX, made education easier to acquire and a more lucrative investment for both black and white women. There have been radical changes in the distribution of education in recent years across the genders, with women now receiving more college degrees than men (Buchmann et al. 2008). As women become more able to turn high levels of education into better employment, and better access to health care and

health-care-related information, we would expect, as indeed we find, an emerging mortality risk gradient across education levels. This gradient emerges among black women in the same cohort as white women, but is significantly less steep. The contrast in the slopes likely reflects the additional discriminatory barriers black women face in the labor market and the increased exposure to discrimination and stigma.

Interestingly, our disaggregated analyses suggest different patterns in gender disparities within race/ethnic groups. While a recent study found that the education-mortality disparity was larger for men than for women (Ross et al. 2012), our results suggest that this finding does not hold across all cohorts, and in fact, the reverse may be true among blacks. Indeed, our results for whites are in line with the results from Zajacova and Hummer (2009), who found no gender differences in education-mortality disparities among younger cohorts. The results for blacks presented here, however, show that the education-mortality gradient for women is steeper than it is for men in the two most recent cohorts. Thus, compared to black men, black women are better able to translate their educational achievement into health benefits.

Our results differ slightly from those of Montez et al. (2012), which suggest a linear decline in mortality risk through 11 years of education, a step change at 12 years, and then continued linear decrease. We find that among white males, 9 years of education is a critical point for changes in the slope. This difference may be due to our examination of a broader range of cohorts, or due to our use of the nonlinear spline and Cox proportional hazard models. While the models fit by Montez et al. are more flexible than most of those used previously in the education-mortality literature, they impose a single slope model for those with 0–11 years of education, while our approach does not. Like Montez et al. (2012) and Rogers et al. (2010), we find that the health returns to education do not diminish across the higher levels of educational achievement.

Our approach of combining high school diploma recipients and GED recipients into a single category may be problematic, as research has shown that there are differences in health returns to education between these two groups (Rogers et al. 2010). Moreover, there may be additional issues with the way that NHIS top codes years of education at 18. It may be that there is a continued decrease in mortality risk beyond 18 years of education that we are unable to capture. Finally, because our analysis is not a true age-period-cohort analysis, some of the cohort differences observed may be due in part to age and period effects. We have reduced these possible confounding factors by restricting our age range to the working age population.

Despite these limitations, this study sheds light on how the nonlinear relationship between education and mortality varies by race/ethnicity and across cohorts. Using cohort rather than age allows us to examine groups of people who received their education during similar historical periods, which have important implications for access to both education and the labor market. Our results suggest that the extent to which fundamental cause theory can explain health disparities may be contingent upon the opportunities available for individuals to capitalize on their SES. More research is needed to understand the processes that differentiate the returns to education between demographic groups, especially the stagnation in the health returns to education in the most recent cohort of black men, and their significantly less steep slope compared to returns for all other demographic groups. Policymakers should continue to address not only the racial/ethnic differences in access to quality education, but also how both structural and interpersonal discrimination may undermine the health benefits of education. Insight into the shape and magnitude of the education-mortality gradient can ultimately increase life expectancies.

References

- Ruppert, D.; Wand, MP.; Carroll, RJ. Semiparametric regression. Boston, MA: Cambridge University Press; 2003.
- Adler NE, Rehkopf DH. U.S. disparities in health: Descriptions, causes, and mechanisms. *Annual Review of Public Health*. 2008; 29:235–252.
- Altonji JG, Blank RM. Race and gender in the labor market. *Handbook of Labor Economics*. 1999; 3:3143–3259.
- Antonovsky A. Social class, life expectancy and overall mortality. *Milbank Memorial Fund Quarterly*. 1967; 45:31–73. [PubMed: 6034566]
- Arrow, K. The theory of discrimination. In: Ashenfelter, O.; Rees, A., editors. *Discrimination in Labor Markets*. Princeton, NJ: Princeton University Press; 1973. p. 3–33.
- Backlund E, Sorlie PD, Johnson NJ. A comparison of the relationships of education and income with mortality: The national longitudinal mortality study. *Social Science & Medicine*. 1999; 49:1373–1384. [PubMed: 10509827]
- Bertrand M, Mullainathan S. Are Emily and Greg more employable than Lakisha and Jamal? A field experiment on labor market discrimination. *American Economic Review*. 2004; 94:991–1013.
- Bonacich E. Advanced capitalism and black/white race relations in the United States: A split labor market interpretation. *American Sociological Review*. 1976; 41:34–51.
- Brunner E. Socioeconomic determinants of health: Stress and the biology of inequality. *British Medical Journal*. 1997; 314:1472. [PubMed: 9167568]
- Buchmann C, DiPrete TA, McDaniel A. Gender inequalities in education. *Annual Review of Sociology*. 2008; 34:319–337.
- Burnham KP, Anderson DR. Multimodel inference. *Sociological Methods and Research*. 2004; 33:261–304.
- Cancian FM, Gordon SL. Changing emotion norms in marriage: Love and anger in US women's magazines since 1900. *Gender and Society*. 1988; 2:308–342.
- Card D, Lemieux T. Can falling supply explain the rising return to college for younger men? A cohort-based analysis. *Quarterly Journal of Economics*. 2011; 116:705–746.
- Donohue JJ III, Heckman J. Continuous versus episodic change: The impact of civil rights policy on the economic status of blacks. *Journal of Economic Literature*. 1991; 29:1603–1643.
- Eilers PHC, Marx BD. Flexible smoothing with B-splines and penalties. *Statistical Science*. 1996; 11:89–102.
- Elo IT, Martikainen P, Smith KP. Socioeconomic differentials in mortality in Finland and the United States: The role of education and income. *European Journal of Population*. 2006; 22:179–203.
- Everett BG, Rogers RG, Hummer RA, Krueger PM. Trends in educational attainment by race/ethnicity, nativity, and sex in the United States, 1989–2005. *Ethnic and Racial Studies*. 2011; 34(9):1543–1566. [PubMed: 22649275]
- Fryer RG Jr, Heaton PS, Levitt SD, Murphy KM. Measuring the impact of crack cocaine. NBER Working Paper. 2005; No. 11318
- Fuchs, VR. *The service economy*. New York: National Bureau of Economic Research; 1968.
- Goldin C. Female labor force participation: The origin of black and white differences, 1870 and 1880. *Journal of Economic History*. 1977; 37:87–108.
- Goldin C, Katz LF. The origins of technology-skill complementarity. *Quarterly Journal of Economics*. 2011; 113:693–732.
- Goldin C, Katz LF. The returns to skill in the United States across the twentieth century. National Bureau of Economic Research Working Paper Series. 1999; No. 7126 <http://www.nber.org/papers/w7126>. Cited March 11, 2011.
- Hanushek EA, Rivkin SG. School quality and the black-white achievement gap. NBER working paper. 2006; No. 12651
- Holzer HJ. The labor market and young black men: Updating Moynihan's perspective. *Annals of the American Academy of Political and Social Science*. 2009; 621:47–69.

- Holzer, HJ.; Raphael, S.; Stoll, MA. Will employers hire former offenders? Employer preferences, background checks, and their determinants. In: Pattillo, M.; Wieman, D.; Wester, B., editors. *Imprisoning America: The social effects of mass incarceration*. New York: Russell Sage Foundation; 2004. p. 205-243.
- Howell DR, Wolff EN. Trends in the growth and distribution of skills in the U.S. workplace, 1960–1985. *Industrial & Labor Relations Review*. 1991; 44:486–502.
- Hummer, RA.; Lariscy, JT. Educational attainment and adult mortality. In: Rogers, RG.; Crimmins, EM., editors. *International handbook of adult mortality*. New York: Springer; 2011. p. 241-261.
- Hurvich CM, Simonoff JS, Tsai CL. Smoothing parameter selection in nonparametric regression using an improved Akaike Information Criterion. *Journal of the Royal Statistical Society: Series B*. 1998; 60:271–293.
- Jacobs JA. Gender inequality and higher education. *Annual Review of Sociology*. 1996; 22:153–185.
- Jemal A, Ward E, Hao Y, Thun M. Mortality from leading causes by education and race in the United States, 2001. *American Journal of Preventive Medicine*. 2008; 34:1–8. [PubMed: 18083444]
- Kitagawa, EM.; Hauser, PM. *Differential mortality in the United States: A study in socioeconomic epidemiology*. Cambridge, MA: Harvard University Press; 1973.
- Kom EL, Graubard BI, Midthune D. Time-to-event analysis of longitudinal follow-up of a survey: Choice of the time-scale. *American Journal of Epidemiology*. 1997; 145:72–80. [PubMed: 8982025]
- Krieger N, Rowley DL, Herman AA, Avery B, Phillips MT. Racism, genderism, and social class: Implications for studies of health, disease, and well-being. *American Journal of Preventive Medicine*. 1993; 9(6):312–323.
- Kunitz SJ. Sex, race and social role—history and the social determinants of health. *International Journal of Epidemiology*. 2007; 36:3–10. [PubMed: 17251243]
- Levy F, Murnane RJ. U.S. earnings levels and earnings inequality: A review of recent trends and proposed explanations. *Journal of Economic Literature*. 1992; 30:1333–1381.
- Link BG, Phelan J. Social conditions as fundamental causes of disease. *Journal of Health and Social Behavior*. 1995;80–94. extra issue.
- Lochner K, Hummer R, Bartee S, Wheatcroft G, Cox C. The public-use National Health Interview Survey Linked Mortality Files: Methods of re-identification risk avoidance and comparative analysis. *American Journal of Epidemiology*. 2008; 168:336–344. [PubMed: 18503037]
- Lynch SM. Explaining life course and cohort variation in the relationship between education and health: The role of income. *Journal of Health and Social Behavior*. 2006; 47:324–338. [PubMed: 17240923]
- Lynch SM. Cohort and life-course patterns in the relationship between education and health: A hierarchical approach. *Demography*. 2003; 40:309–331. [PubMed: 12846134]
- Masters RK, Hummer RA, Powers DA. Educational differences in U.S. adult mortality: A cohort perspective. *American Sociological Review*. 2012; 77:548–572.
- Meara ER, Richards S, Cutler DM. The gap gets bigger: Changes in mortality and life expectancy, by education, 1981–2000. *Health Affairs*. 2008; 27:350–360. [PubMed: 18332489]
- Miech R, Pampel F, Kim J, Rogers RG. The enduring association between education and mortality: The role of widening and narrowing disparities. *American Sociological Review*. 2011; 76:913–934.
- Mirowsky J, Ross CE. Education, personal control, lifestyle and health: A human capital hypothesis. *Research on Aging*. 1998; 20(4):415–449.
- Mirowsky, J.; Ross, CE. *Education, social status, and health*. New York: Aldine de Gruyter; 2003.
- Molla MT, Madans JH, Wagener DK. Differentials in adult mortality and activity limitation by years of education in the United States at the end of the 1990s. *Population and Development Review*. 2004; 30:625–646.
- Montez JK, Hayward MD, Brown DC, Hummer RA. Why is the educational gradient of mortality steeper for men? *Journals of Gerontology Series B: Psychological Sciences and Social Sciences*. 2009; 64B:625–634.

- Montez JK, Hummer RA, Hayward M. Educational attainment and adult mortality in the United States: A systematic analysis of functional form. *Demography*. 2012; 49(1):315–336. [PubMed: 22246797]
- National Center for Atmospheric Research (NCAR). Verification: Forecast verification utilities. R package version 1.34. 2012. <http://CRAN.R-project.org/package=verification>
- National Center for Health Statistics. National Health Interview Surveys: Core Data. Hyattsville, MD: U.S. Department of Health and Human Services; 1997–2000. Various years.
- National Center for Health Statistics. The 1986–2000 National Health Interview Survey Linked Mortality Files: Matching Methodology. Hyattsville, Maryland: 2005. Available at http://www.cdc.gov/nchs/data/datalinkage/matching_methodology_nhis_final.pdf
- National Center for Health Statistics. National Health Interview Survey, Linked Mortality Public-Use Data File: 1986–2006 Survey Years. Hyattsville, MD: U.S. Department of Health and Human Services; 2010.
- Neal D. Why has black-white skill convergence stopped? *Handbook of the Economics of Education*. 2006; 1:511–576.
- OECD. Employment outlook. Paris: OECD; 2000.
- Olshansky SJ, Antonucci T, Berkman L, Binstock RH, Boersch-Supan A, Cacioppo JT, Carnes BA, Carstensen LL, Fried LP, Goldman DP, Jackson J, Kohli M, Rother J, Zheng Y, Rowe J. Differences in life expectancy due to race and educational differences are widening, and many may not catch up. *Health Affairs*. 2012; 31:1803–1813. [PubMed: 22869659]
- Pappas G, Queen S, Hadden W, Fisher G. The increasing disparity in mortality between socioeconomic groups in the United States, 1960 and 1986. *New England Journal of Medicine*. 1993; 329:103–109. [PubMed: 8510686]
- Phelan JC, Link BG, Diez-Roux A, Kawachi I, Levin B. Fundamental causes of social inequalities in mortality: A test of the theory. *Journal of Health and Social Behavior*. 2004; 45:265–285. [PubMed: 15595507]
- Phelps ES. The statistical theory of racism and sexism. *American Economic Review*. 1972; 62:659–661.
- Preston SH, Elo IT. Are educational differentials in adult mortality increasing in the United States? *Journal of Aging and Health*. 1995; 7:476–496. [PubMed: 10165966]
- Ramsey J, Ripley B. Pspine: Penalized smoothing splines. R package version. 2007:1–0.
- Rehkopf D, Berkman LF, Coull B, Krieger N. The non-linear risk of mortality by income level in a healthy population: US National Health and Nutrition Examination Survey mortality follow-up cohort, 1988–2001. *BMC Public Health*. 2008; 8:383. [PubMed: 19000311]
- Rehkopf DH, Krieger NH, Coull B, Berkman LF. Biologic risk markers for coronary heart disease: Nonlinear associations with income. *Epidemiology*. 2010; 21:38–46. [PubMed: 20010209]
- Rehkopf DH, Haughton LT, Chen JT, Waterman PD, Subramanian SV, Krieger N. Monitoring socioeconomic disparities in death: Comparing individual-level education and area-based socioeconomic measures. *American Journal of Public Health*. 2006; 96:2135–2138. [PubMed: 16809582]
- Rogers RG, Hummer RA, Nam, CB. Living and dying in the USA: Behavioral, health, and social differences in adult mortality. San Diego, CA: Academic Press; 2000.
- Rogers RG, Everett BG, Zajacova A, Hummer RA. Educational degrees and adult mortality risk in the United States. *Biodemography and Social Biology*. 2010; 56:80–99. [PubMed: 20589989]
- Rogers RG, Hummer RA, Everett, BG. Educational differentials in US adult mortality: An examination of mediating factors. *Social Science Research*. 2012. doi: <http://dx.doi.org/10.1016/j.ssresearch.2012.09.003>
- Ross CE, Wu CL. The links between education and health. *American Sociological Review*. 1995; 60(5):719–745.
- Ross CE, Masters RK, Hummer RA. Education and the gender gaps in health and mortality. *Demography*. 2012; 49:1157–1183. [PubMed: 22886759]
- Ruppert, D.; Wand, MP.; Carroll, RJ. Semiparametric regression. Boston, MA: Cambridge University Press; 2003.

- Ryan AM, Gee GC, Laflamme DF. The association between self-reported discrimination, physical health, and blood pressure: Findings from African Americans, black immigrants, and Latino immigrants in New Hampshire. *Journal of Health Care for the Poor and Underserved*. 2006; 17:116–132. [PubMed: 16809879]
- Steenland K, Henley J, Thun M. All-cause and cause-specific death rates by educational status for two million people in two American Cancer Society cohorts, 1959–1996. *American Journal of Epidemiology*. 2002; 156:11–21. [PubMed: 12076884]
- Williams DR, Neighbors HW, Jackson JS. Racial/ethnic discrimination and health: Findings from community studies. *American Journal of Public Health*. 2008; 98:S29–37. [PubMed: 18687616]
- Wood, SN. *Generalized additive models: An introduction with R*. CRC Press; Boca Raton, FL: Taylor & Francis Group: 2006.
- Zajacova A. Education, gender, and mortality: Does schooling have the same effect on mortality for men and women in the US? *Social Science & Medicine*. 2006; 63:2176–2190. [PubMed: 16781036]
- Zajacova A, Hummer RA. Gender differences in education effects on all-cause mortality for white and black adults in the United States. *Social Science & Medicine*. 2009; 69:529–537. [PubMed: 19589633]

Figure 1. Males

Figure 2. Females

Table 1
Descriptive Information on Mortality and Education by Sex, Race/Ethnicity, and Cohort

	Non Hispanic White Males						Non-Hispanic Black Males					
	Born						Born					
	1921-1969	1921-1936	1937-1953	1954-1969	1921-1969	1921-1936	1937-1953	1954-1969	1921-1969	1921-1936	1937-1953	1954-1969
Yrs of Education	%	%	%	%	%	%	%	%	%	%	%	%
0 to 8	4.39	11.96	3.80	1.49	30.11	30.11	8.50	2.47	30.11	30.11	8.50	2.47
9 to 11	7.68	12.70	6.95	6.35	21.91	21.91	16.49	11.18	21.91	21.91	16.49	11.18
12	37.93	37.34	36.44	40.80	29.33	29.33	41.61	48.55	29.33	29.33	41.61	48.55
13 to 15	19.45	13.93	20.08	21.24	10.35	10.35	18.39	22.68	10.35	10.35	18.39	22.68
16	16.81	12.11	16.94	18.20	3.80	3.80	8.47	10.62	3.80	3.80	8.47	10.62
17 to 18	13.74	11.96	15.79	11.13	4.50	4.50	6.55	4.50	4.50	4.50	6.55	4.50
% Dead by Education												
0 to 8	29.41	47.15	17.96	6.21	35.29	51.91	23.11	10.27	35.29	51.91	23.11	10.27
9 to 11	18.84	42.02	14.00	4.02	21.35	46.40	18.63	6.36	21.35	46.40	18.63	6.36
12	10.37	31.85	9.05	2.23	12.16	38.26	13.41	3.93	12.16	38.26	13.41	3.93
13 to 15	8.27	33.67	6.94	1.83	10.15	33.10	11.67	3.94	10.15	33.10	11.67	3.94
16	5.91	24.82	4.99	1.09	7.72	32.62	9.63	1.72	7.72	32.62	9.63	1.72
17 to 18	5.98	21.56	4.43	1.09	8.13	27.12	7.06	2.33	8.13	27.12	7.06	2.33
N	224,018	42,257	123,447	58,314	32,722	6,174	17,526	9,022	32,722	6,174	17,526	9,022
N Died	25,885	14,187	10,397	1,301	5,743	2,720	2,615	408	5,743	2,720	2,615	408
	Non-Hispanic White Females						Non-Hispanic Black Females					
	Born						Born					
	1921-1969	1921-1936	1937-1953	1954-1969	1921-1969	1921-1936	1937-1953	1954-1969	1921-1969	1921-1936	1937-1953	1954-1969
Yrs of Education												
0 to 8	3.25	8.66	2.72	1.18	6.86	22.21	5.92	1.68	6.86	22.21	5.92	1.68
9 to 11	7.74	14.07	7.32	4.94	16.28	25.26	17.04	11.34	16.28	25.26	17.04	11.34
12	43.64	49.68	44.08	39.49	42.03	33.56	42.97	44.28	42.03	33.56	42.97	44.28
13 to 15	21.09	14.83	20.96	24.82	19.98	9.75	19.14	25.57	19.98	9.75	19.14	25.57

	Non Hispanic White Males					Non-Hispanic Black Males				
	Born					Born				
	1921-1969	1921-1936	1937-1953	1954-1969		1921-1969	1921-1936	1937-1953	1954-1969	
16	14.33	7.39	13.49	19.68		8.86	4.68	8.09	11.77	
17 to 18	9.95	5.37	11.44	9.91		6.00	4.55	6.85	5.36	
% Dead by Education										
0 to 8	22.53	36.23	13.17	4.11		26.10	37.20	17.87	6.08	
9 to 11	15.80	31.65	11.05	2.90		16.27	34.34	13.57	5.06	
12	7.64	22.24	5.48	1.60		8.55	25.77	8.84	2.87	
13 to 15	5.09	21.41	4.19	0.97		5.93	25.97	6.77	1.75	
16	3.28	17.57	2.93	0.71		4.76	19.03	5.13	1.96	
17 to 18	3.33	15.05	2.88	0.68		4.96	18.62	4.57	0.75	
N	237,014	46,560	128,810	61,644		46,016	8,663	24,435	12,918	
N Died	19,277	11,282	7,127	868		5,527	2,682	2,465	380	

Source: Derived from NHIS-LMF.

Fit Statistics for Cox Proportional Hazard Models

Table 2

	Cox PH: Spline	Cox PH: Spline with GEE	Cox, PH: Spline, Frailty with Gamma Distribution	Cox, PH: Spline, Frailty with Gaussian Distribution	Cox PH: Linear	Cox PH: Categorical (Less than HS, HS Graduate)	Cox PH: Spline	Cox PH: Spline with GEE	Cox PH: Spline, Frailty with Gaussian Distribution	Cox, PH: Spline, Frailty with Gaussian Distribution	Cox PH: Linear	Cox PH: Categorical (Less than HS, HS Graduate)
Non-Hispanic White Males												
Cohort 1, Born 1921 to 1936 (N=42,257; Dead=14,178)												
AIC	154748	281365	281365	281326	154779	154910	125450	225969	225969	225937	12547	126340
ROC Area	0.6	0.618	0.619	0.622	0.599	0.593	0.586	0.602	0.603	0.605	0.589	0.583
Passed Schoenfeld Test	Yes	Yes	Yes	Yes	No	No	Yes	No	Yes	Yes	No	No
Cohort 2, Born 1937 to 1953 (N=123,447; Dead=10,397)												
AIC	107192	213871	213841	213847	107208	107368	74048	147202	147203	147160	74084	74065
ROC Area	0.617	0.642	0.648	0.645	0.616	0.612	0.619	0.642	0.643	0.651	0.617	0.616
Passed Schoenfeld Test	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No
Cohort 3, Born 1954 to 1969 (N=58,314; Dead=1,301)												
AIC	9539	23797	23797	23797	9540	9559	6628	16344	16344	16344	6629	6629
ROC Area	0.641	0.673	0.676	0.676	0.642	0.634	0.634	0.665	0.669	0.669	0.633	0.633
Passed Schoenfeld Test	Yes	Yes	Yes	Yes	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Non-Hispanic Black Males												
Cohort 1, Born 1921 to 1936 (N=6,174; Dead=2,720)												
AIC	21487	43124	43044	43055	21499	21482	23109	44168	4107	44122	23124	23117
ROC Area	0.595	0.615	0.648	0.638	0.595	0.596	0.576	0.595	0.619	0.614	0.569	0.569
Passed Schoenfeld Test	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	Yes	Yes	Yes
Cohort 2, Born 1937 to 1953 (N=17,523; Dead=2,615)												
AIC	20098	44781	44781	44760	20107	20114	20094	42273	4225	42256	20104	20103
ROC Area	0.582	0.622	0.623	0.633	0.584	0.584	0.603	0.634	0.645	0.64	0.601	0.604
Passed Schoenfeld Test	Yes	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cohort 3, Born 1954 to 1969 (N=12,918; Dead=380)												

	Cox PH: Spline	Cox PH: Spline with GEE	Cox, PH: Spline, Frailty with Gamma Distribution	Cox, PH: Spline, Frailty with Gaussian Distribution	Cox PH: Linear	Cox PH: Categorical (Less than HS, HS Graduate)	Cox PH: Spline	Cox PH: Spline with GEE	Cox PH: Spline, Frailty with Gamma Distribution	Cox, PH: Spline, Frailty with Gaussian Distribution	Cox PH: Linear	Cox PH: Categorical (Less than HS, HS Graduate)
ROC Area	0.612	0.657	0.662	0.678	0.612	0.609	0.604	0.669	0.673	0.711	0.604	0.607
Passed Schoenfeld Test	Yes	No	Yes	Yes	Yes	No	Yes	Yes	No	No	Yes	Yes

Source: National Health Interview Survey-Linked Mortality Files.

Notes: Models control for year of survey, age in the time scale, and survey Change; PH=Proportional Hazards

Table 3
Predicted Hazard Ratios by Education Level from Nonlinear Cox Proportional Hazard Models

Years of Education	Non-Hispanic White Males				Non-Hispanic Black Males			
	Born	Born	Born	Born	Born	Born	Born	Born
	1921-1936	1937-1953	1954-1969		1921-1936	1937-1953	1954-1969	
0	1.808	2.132	5.529		1.278	1.652	2.354	
1	1.784	1.970	4.807		1.256	1.458	2.435	
2	1.740	1.855	6.424		1.245	1.752	3.254	
3	1.725	1.806	4.953		1.217	1.510	2.507	
4	1.692	1.725	4.855		1.220	1.459	2.858	
5	1.637	1.654	3.935		1.210	1.425	2.514	
6	1.575	1.571	3.525		1.202	1.402	2.273	
7	1.499	1.495	3.127		1.182	1.365	2.404	
8	1.422	1.415	2.746		1.157	1.331	2.028	
9	1.328	1.313	2.404		1.119	1.289	1.828	
10	1.230	1.203	1.952		1.079	1.220	1.498	
11	1.141	1.092	1.579		1.017	1.143	1.270	
12	1.045	0.990	1.298		0.942	1.040	1.104	
13	0.992	0.917	1.446		0.893	1.171	1.260	
14	0.913	0.859	0.850		1.220	0.826	0.810	
15	0.859	0.811	0.945		0.767	0.981	1.000	
16	0.771	0.780	0.680		0.705	0.730	0.662	
17	0.714	0.753	0.805		0.679	0.784	0.748	
18	0.647	0.733	0.701		0.628	0.682	0.626	
Years of Education	Non-Hispanic White Females				Non-Hispanic Black Females			
	Born	Born	Born	Born	Born	Born	Born	Born
	1921-1936	1937-1953	1954-1969		1921-1936	1937-1953	1954-1969	
0	2.145	3.597	5.755		1.237	2.042	3.651	

	Non-Hispanic White Males				Non-Hispanic Black Males			
	Born		Born		Born		Born	
	1921-1936	1937-1953	1954-1969	1970-1983	1921-1936	1937-1953	1954-1969	1970-1983
1	2.016	2.858	5.104		1.235	1.614	3.065	
2	1.919	3.158	5.053		1.257	1.865	3.781	
3	1.868	2.945	4.306		1.266	1.732	3.387	
4	1.759	2.915	4.015		1.266	1.763	4.204	
5	1.704	2.974	4.137		1.259	1.774	3.655	
6	1.619	2.560	3.633		1.247	1.647	2.995	
7	1.537	2.401	3.287		1.225	1.640	2.942	
8	1.452	2.145	2.915		1.198	1.553	2.504	
9	1.340	1.944	2.535		1.165	1.487	2.282	
10	1.224	1.662	2.117		1.111	1.349	1.835	
11	1.103	1.383	1.699		1.042	1.190	1.449	
12	0.996	1.149	1.363		0.969	1.062	1.179	
13	0.924	1.179	1.413		0.899	1.105	1.254	
14	0.850	0.789	0.829		0.820	0.745	0.753	
15	0.808	0.946	0.988		0.755	0.878	0.607	
16	0.750	0.712	0.651		0.682	0.656	0.613	
17	0.720	0.803	0.766		0.623	0.705	0.696	
18	0.674	0.743	0.676		0.561	0.641	0.576	

Source: National Health Interview Survey- Linked Mortality Files.

Models control for year of survey, age in the time scale, and survey change

Table 4
Hazard Ratios for Education Interaction Terms Derived from Hazard Models to Assess Significant Differences in Nonlinear Mortality curves

	Cohorts					
	1921-1936		1937-1953		1954-1969	
	HR	P	HR	P	HR	P
Black Males*Education (White males)	1.02	0.00	1.05	0.00	1.04	0.00
Black Females* Education (White females)	1.03	0.00	1.03	0.00	1.01	0.00
White Males* Education (White females)	1.01	0.10	1.01	0.77	1.01	0.05
Black Males* Education (Black females)	1.01	0.40	1.03	0.01	1.01	0.70
Demographic Subgroups						
	Whites Males	Black Males	White Females	Black Females		
Born 1937-1953* Education (Born 1921-1936)	0.93	0.97	0.70	0.96	0.00	0.03
Born 1954-1969* Education (Born 1937-1953)	0.96	0.95	0.01	0.96	0.00	0.11

Source: NHIS-LMF: Referent is in parentheses; HR=Hazard Ratio; P=p value

Notes: All models control for age, year, and survey change; where appropriate, models control for race/ethnicity, sex, or cohort: HR=Hazard Ratio