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Gender Differences in Education Effects on All-Cause Mortality for White and Black Adults in the United States

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Abstract

The existence of education differentials in adult mortality has been well established. The issue of gender differences in the education-mortality association, however, remains an open question, despite its importance for understanding of causal pathways through which education affects health outcomes. The goal of this paper is to analyze gender differences in education gradients in mortality among non-Hispanic white and black U.S. adults born between 1906 and 1965. The analysis is based on data from the 1986–2000 National Health Interview Surveys linked to the National Death Index through 2002 (NHIS-LMF) with over 700,000 respondents. Full-sample and cohort-stratified Cox proportional hazard models of all-cause mortality were estimated. Results indicate a great deal of similarity between men and women in the education-mortality association, with some exceptions. The most notable difference is the steeper educational gradient at high schooling levels for white men compared to white women. This difference was fully explained by marital status. No systematic gender differences in the relationship between education and adult mortality were observed among black adults in any birth cohorts. The findings suggest that men do not benefit from educational attainment uniformly more than women.

Keywords

USA; white adults; education; gender; mortality; marital status; black adults

The strong effects¹ of education on adult mortality have been described exhaustively (Kitagawa & Hauser, 1973; Lleras Muney, 2005; Pappas, Queen, Hadden, & Fisher, 1993; Preston & Elo, 1995; Sorlie, Backlund, & Keller, 1995). It has long been known that the education effects may vary across different population subgroups (Bassuk, Berkman, & Amick, 2002; House, Kessler, Herzog, Mero, Kinney, & Breslow, 1990; Lauderdale, 2001; Lynch, 2006). Gender differences are of particular interest because the comparison of the education-mortality association for men

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¹Most of the studies cited here did not formally estimate causal effects of education on mortality but rather describe their association. Although we sometimes use causal language in the text, we similarly do not purport to determine the causal links between education and mortality. Nonetheless, there is emerging literature that shows that a substantial part of the association between education and mortality is due to the causal effects of schooling (Chandola, Clarke, Morris, & Blane, 2008; Glied & Lleras-Muney, 2008; Lleras Muney, 2005). Hence, while we cannot make strong causal claims on the basis of our analysis, we suggest that the observed association between education and mortality may be interpreted as at least partly *caused* by educational attainment.

and women can be combined with our knowledge about socioeconomic returns to education to gain insights into the causal pathways through which education impacts health outcomes. In Europe, studies suggested that education gradients in mortality are larger for men than for women (i.e., Mackenbach, Kunst, Groenhouf, Borgan, & Costa, 1999). We lack evidence-supported consensus, however, on how educational attainment relates to men's and women's risks of dying among American adults.

Documenting the interaction of education and gender, two critical dimensions of the social stratification matrix, is critical for disentangling the causal pathways from social status to health. We need to understand the consequences of educational inequalities for men and women in health and longevity, attributes that are demonstrably valuable to individuals and societies. The goal of this paper is to provide a detailed examination of gender differences in the association between education and all-cause mortality in the United States, focusing on non-Hispanic white and black adults born between 1906 and 1965. We show that the gross relative educational inequalities in all-cause mortality are generally comparable for men and women, with an important exception at the highest education levels for some demographic groups.

PREVIOUS RESEARCH

The existing body of health literature that focused on the interplay of gender and education offers a somewhat inconsistent picture (Macintyre & Hunt, 1997; Matthews, Manor, & Power, 1999). Findings from European countries typically show steeper education gradients, or relative mortality differentials, for men than for women. The gender differences are observed particularly at working ages (Elo, Martikainen, & Smith, 2006; Koskinen & Martelin, 1994; Mustard & Etches, 2003; Valkonen, 1989), while at older ages the gender differences are typically smaller (Huisman, Kunst, Andersen, Bopp, & al., 2004; Huisman, Kunst, Bopp, & al., 2005; Mackenbach et al., 1999). It is important to note that these studies did not conduct formal tests of the statistical significance of the observed gender differences in the education-mortality association – the summary above comes from an examination of mortality ratios and their confidence intervals estimated separately for men and for women. Such examination, however, is not equivalent to proper statistical tests of the difference (Austin & Hux, 2002; Schenker & Gentleman, 2001).

The findings from European contexts cannot be automatically extended to U.S. adults. Considerable cross-country variation exists in both relative and absolute education differentials in mortality for men and women (Huisman et al., 2004; Valkonen, 1989). Recently, Kohler, Martikainen, Smith, and Elo (2008), focusing on both relative and absolute mortality differentials, showed a clear variation in education gradients within Europe and between European countries and the United States. This variation suggests large differences in how a given schooling level translates to other resources and stratification attributes, including mortality levels, across national contexts (Card, 1999). The structure of gender and socioeconomic inequalities in the United States differs substantially from European countries. Accordingly, the patterns observed for the gender and education interaction in longevity may be different in the United States, compared to Europe.

While a number of studies have described larger education differentials in mortality for U.S. men than for women (Bassuk et al., 2002; Feldman, Makuc, Kleinman, & Cornoni-Huntley, 1989; Mackenbach et al., 1999; Montez, Hayward, Brown, & Hummer, 2009; Pappas et al., 1993), other researchers reported little or no systematic gender differentials (Elo et al., 2006; Feldman et al., 1989; Kohler et al., 2008; Sorlie et al., 1995). Finally, some U.S. analyses found slightly larger relative educational differences in mortality for women than men (Deaton & Paxson, 1999), including the seminal work by Kitagawa and Hauser (1973) using data from 1960. Some of these inconsistencies may be due to differences in the samples --- ages, periods,

and cohorts studied, or to different analytic approaches. Most relevant, however, is that none of these studies tested whether the observed gender differences in the education-mortality relationship were statistically significant or occurred by chance.

We found only four papers that explicitly focused on analyzing the interaction between gender and education on adult mortality in the United States. Christenson and Johnson (1995), using Michigan death certificate and census data, reported substantively small but statistically significant gender differences in the education patterns, in the opposite direction at different education levels: mortality differences between postsecondary versus secondary schooling levels were larger for men, while differences between presecondary versus secondary levels were larger for women. Two other studies were based on nationally representative data, albeit with relatively small sample sizes, and employed a parsimonious, but not necessarily the most appropriate, continuous specification of education. Both reported no significant gender difference in the effect of education on all-cause mortality (McDonough, Williams, House, & Duncan, 1999; Zajacova, 2006). Finally, a recent paper by Montez et al. (2009) found larger education gradients in all-cause mortality for men, but only among unmarried adults. Hence, we still lack a strong consensus on how gender and education interact to affect mortality patterns among American adults.

When analyzing educational gradients in mortality, an important variable to consider is marital status. Marital status has been noted as one of the major factors that interacts with socioeconomic status to affect health. A classic study by Smith and Waitzman (1994) demonstrated the “double jeopardy” effect of low economic status and unmarried status, especially for men. More recently, Kohler et al (2008) showed how also educational effects on mortality depend on men’s and women’s marital status, and Montez et al. (2009) found that for white American adults, marital status played a crucial role in understanding gender differentials in the education-mortality association.

This study extends previous work in a number of important directions. The main contribution is in providing a detailed description of how gender moderates educational gradients in U.S. adult mortality, including formal statistical tests for the moderating effect of gender. The dataset we employ is considered the authoritative source for health information in the United States, and includes excellent mortality followup. Thanks to the large number of observations (almost 7 million person-years of data), we are able to study the *shape* of the education-mortality association, rather than use an a priori specification of the education variable. We examine different birth cohorts separately to avoid possible confounding of the results by strong cohort-specific influences. We include marital status in our analyses as a factor previously shown to play a decisive role when considering how men’s and women’s health benefits from educational attainment differ. And finally, we study black and white adults separately, in contrast to previous studies which either focused exclusively on white adults or controlled for race. There is a strong theoretical motivation for estimating race-stratified analyses when focusing on the education-mortality association. Black adults have significantly lower average education than white adults (Stoops, 2004) and experience higher mortality rates (Arias, 2006). Further, there may be significant race differences in the association of education and mortality (Christenson & Johnson, 1995), which could confound and bias results that are not stratified by race.

MATERIALS AND METHODS

Data

We use the National Health Interview Survey (NHIS) data pooled from 1986–2000, linked to the National Death Index information on all deaths through 2002. The linked data file is called the National Health Interview Survey-Linked Mortality File (NHIS-LMF). The NHIS is an

annual multiple-purpose health survey conducted through face-to-face household interviews by the National Center for Health Statistics (NCHS). NHIS uses a complex multistage stratified sampling design to obtain a sample representative of the civilian non-institutionalized U.S. population. The total household response rate was around 90%; the rates were over 95% in the earlier years of the data used here (Massey, Moore, Parsons, & Tadros, 1989) and declined to 89% in 2000 (CDC, 2002). Additional information about the sampling design is available in Massey et al. (1989) for years 1985–1994 and in Botman et al. (2000) for 1995–2000.

Mortality followup for adults who participated in 1986–2000 NHIS was obtained by linking their records to the National Death Index. Details about the probabilistic matching algorithm have been published elsewhere (Lochner, Hummer, Bartee, Wheatcroft, & Cox, 2008; NCHS, 2005). Ingram et al. (2008) conducted a thorough evaluation of the match quality by comparing the age-sex-race-specific cumulative survival of the NHIS-LMF individuals to that of the U.S. population using U.S. life tables. The NHIS-LMF survey cohorts had slightly higher survival probabilities than official U.S. life table data because the NHIS sampled only non-institutionalized adults. The differences were larger in absolute terms, but less likely to be statistically significant, for black adults compared to white adults (see figures 1–9 and tables 20–23 in Ingram et al., 2008). The comparison showed that for the sample used in the present analysis, the linkage is of very high quality and allows generalizing our findings to the U.S. adult civilian, non-institutionalized population.

Sample

We included non-Hispanic white and non-Hispanic black respondents born between 1906 and 1965 who were 25–80 years old at the time of interview. Respondents from other racial/ethnic categories were excluded because of their substantial heterogeneity and relatively small number of deaths. The lower age limit was selected as an age by which most individuals have completed their schooling. The upper age limit was set because the linkage of the NHIS records to the NDI was less successful for women over 80 compared to men (see Ingram et al. 2008), which could lead to biased findings. The data had no missing values on age, sex, race, and region of residence. Individuals with missing values on education (0.6 percent) were excluded from the analyses.

Measures

Up to 1996, NHIS measured education as the highest completed year of schooling ranging from 0 to 18. In 1997–2000, education was measured in years of schooling up to the 12th year and in secondary and postsecondary degrees at higher levels. We converted the degrees into years of schooling as follows: some college but no degree=13 years, associate's degree=14 years, bachelor's degree=16 years, master's degree=18 years, and professional or doctoral degree=20 years. We employed different specifications of the education variable in different parts of the analysis: as a continuous predictor centered around the mean of 12 years; as a 6-category predictor; and as a single-year dummy predictor for results shown in Figure 1. The reference category in both categorical classifications was 12 years.

Control variables in all regression models included age at interview centered on a model-specific mean, gender (male as reference), region of residence (Northeast as reference), and year of interview (centered on the mean year 1993). Some models also controlled for marital status, coded as married (reference), divorced, widowed, never married, and unknown marital status. All models were estimated separately for white and black adults.

Analysis

We used Cox proportional hazard models to calculate mortality hazard ratios by education. The Cox model is the most widely used approach for analyzing event history data because it

accommodates binary outcomes (e.g., died versus did not die), as well as time to death or censoring at the end of the observation period in 2002. Further, it does not require any specifications for the distribution of the baseline hazard. We also examined alternative estimation approaches, including Poisson regression models with log-duration offset and logistic regression models using a person-year data file structure. Results from these different estimation strategies were comparable to the Cox models and are available on request.

We estimated both sex-stratified models (for Figure 1) and sex-combined models that included education-by-female interaction terms. The statistical significance of the interaction coefficient (s) was used to formally test for sex differences in the effects of education on the mortality hazard. The interaction models also included an age-by-female term in order to account for the different age profile of the mortality hazard for U.S. men and women (Arias, 2006). Some analyses were conducted by birth cohort: 1906–1925, 1926–1945, and 1946–1965. We chose these three cohort groups as a balance between more detailed cohort analyses and parsimony, loosely following Carlson's (2008) 'Good Warriors,' 'Lucky Few,' and 'Baby Boomers' cohorts. We conducted sensitivity analyses using more cohort groups and different threshold years, with no substantive changes to the findings. All analyses were stratified by race and controlled for age at interview, year of interview, and region of residence. The analyses adjusted for sampling design by using survey weights and Taylor-linearized variance estimation available in Stata 10 (StataCorp, 2007).

We completed a number of diagnostic tests to confirm the validity of our regression findings. To test the proportionality of hazards, we conducted three sets of auxiliary analyses. First, we plotted Kaplan-Meier curves and log-log survival plots by sex, categorized education, and cohort. There appeared to be only a slight narrowing of mortality hazards by education with longer follow-up time. Second, we tested for non-zero slopes using scaled Schoenfeld residuals. We plotted the predictor-specific residuals against followup time and also conducted a global statistical test. Both approaches showed no significant violations of the proportionality assumption. And finally, we estimated models where education was allowed to vary with followup time and with age. The results showed that the effect of education weakened slightly with time for the oldest cohort but remained stable for the youngest cohorts among both men and women.

We conducted additional validity checks. We estimated parametric survival models with frailty using a Gompertz-distributed baseline hazard, specifying the random error to follow a Gamma or Inverse Gaussian distribution. We found evidence of a non-zero distributed frailty term but the coefficients for education and the education-by-sex interaction remained substantively unchanged. The results were robust to different methods for dealing with ties in the Cox models, including Breslow, Efron, and exact methods, and to different specifications of the followup time interval: single calendar years or three-month time increments. These tests showed that the main findings remained essentially unchanged under different model assumptions and thus corroborated the validity of our findings.

RESULTS

Table 1 summarizes characteristics of the sample at interview and during followup. Among both white and black adults, there is a statistically significant sex difference in the distribution of all key variables. With regard to education, men are more likely to report the lowest (0–8 years) and highest (16 or more years) attainment levels than women, more of whom report 9–15 years of schooling.

Findings from multivariate proportional hazard models of all-cause mortality by education for men and women are summarized in Tables 2–4, using a continuous specification for the effect

of education (Table 2) and a categorical specification (Tables 3 and 4). The continuous measure is preferred for parsimony while the categorical specification allows a detailed examination of whether sex differences exist at specific education levels, which the continuous measure might obscure.

Among white adults (upper part of Table 2), the effect of education for men in the aggregate sample (column titled “all cohorts”) shows a 6% lower mortality hazard for each additional year of schooling. This effect is slightly larger than has been reported previously using older U.S. data (Elo & Preston, 1996; Zajacova, 2006). The main education effect is smaller among the oldest birth cohorts (4% lower hazard of dying among men born 1906–1925) and larger for younger cohorts (11% lower hazard for each additional year of education among men born 1946–1965). Both the comparison with previous studies and the cohort patterns observed here are consistent with the observation that education is becoming increasingly more important for U.S. health and mortality in recent time periods and/or among younger birth cohorts (Lauderdale, 2001; Lynch, 2003, 2006).

The education-by-female interaction terms are the main coefficients of interest because they offer a formal test of whether men and women differ significantly in the education-mortality association. Among white adults, the significant interaction term indicates that the effect of education is weaker for women than for men. The effect of schooling for women is calculated by exponentiating the sum of the coefficients (log-hazard ratios) for the main effect and the interaction effect: women evidence about 5% lower risk of death for each additional year of education, compared to 6% lower risk among men. Disaggregated by birth cohort, the models suggest that the statistically weaker effect of education for women is driven by the oldest birth cohort. For white adults born after 1925, gender difference in the effect of education is not significantly different from zero.

For the full sample of black adults (lower part of Table 2), the effect of education for men shows about 4% lower risk of dying for each additional year of education. As an important side note, this effect is smaller than among whites, suggesting a flatter education gradient in mortality among black men relative to white men. The difference in the mortality slopes by race is also statistically significant (results not shown), corroborating previous studies (Christenson & Johnson, 1995) and highlighting the importance of estimating race-stratified models. The cohort patterns in the effect of schooling are similar to those observed among whites: the effect is weakest among the oldest men (a 2% lower mortality hazard for each additional year of schooling among the 1906–25 birth cohort) and stronger for younger men (a 9% lower hazard of mortality for each additional year of education for the 1946–65 birth cohort).

The education-by-female interaction term for the black sample is significant neither in the “all-cohorts” model nor in any of the three cohort groups, indicating that black men and women exhibit the same statistical association between education and the mortality hazard.

Table 3 presents results using education categorized into 6 levels, in order to evaluate gender differences separately at various points along the educational attainment range. Among whites (upper part of Table 3), the effect of education for men is graded, persistent and strong across all schooling levels and across all birth cohorts. The mortality hazards are all significantly different from the reference category (12 years) at $p < .01$, with a single exception for men with 13–15 years of schooling born in 1926–45. In the aggregate sample, for instance, white men with 13–15 years of schooling have an 8% lower risk of dying during the follow-up period compared to men who completed 12 years of education; those with 16 years of schooling have 28% lower risk; and those with more than 16 years have a 37% lower risk of dying. As with

the linear specification, the education gradient is clearly stronger among younger birth cohorts than older birth cohorts.

Most of the education-by-female interaction coefficients do not attain statistical significance for whites. That is, at the pre-secondary levels and for postsecondary education up to 15 years, the relationship between education and mortality risk for men and women is not significantly different either for the aggregate sample or for the specific birth cohorts (with one exception, 9–11 years relative to 12 years for birth cohort 1926–45, where the mortality differential is significantly *stronger* for women than men). However, at 16 years and 17+ years of schooling, corresponding roughly to a bachelor's degree and more advanced degrees, respectively, the effects of education are weaker for women compared to men. This gender difference in the effect of education is statistically significant in the aggregate sample and in the 1926–45 birth cohort, although not in the oldest or the youngest cohort groups.

Among black men (lower part of Table 3), education effects across most schooling levels appear weaker than among white men. However, the findings still indicate a clear and highly significant mortality gradient, especially among the younger two birth cohorts. Black men with more than 16 years of schooling experience between 12% and 62% lower risks of dying compared to men with 12 years of schooling, depending on the birth cohort.

The education-by-female coefficients for black adults are consistent across all three groups of birth cohorts and in the aggregate sample: none of the 20 interaction terms is statistically significant. This null finding corroborates results from the linear-specification models shown in Table 2, which also describes no observable gender differences in the association between education and mortality risk among black men and women.

Figure 1 shows graphically the shape of the education-mortality association for men and women. The Figure displays results from models for men and women, stratified by race and birth cohort, using dummies for each year of schooling (with 12 years as reference). The plots reveal in detail the similarity between men and women in the mortality patterns by education across the different birth cohorts and racial groups: the hazard ratios for men and women are close in all six plots; the confidence intervals overlap for most of the educational attainment range. The Figure also emphasizes an absence of severe departures from linearity in the effects of education on the mortality hazard, as well as the overall pattern of stronger education effects for younger cohorts relative to older cohorts.

In the final set of models we tested previous researchers' observation (Kohler et al., 2008; Smith & Waitzman, 1994) of a strong 'synergy' between education and marital status on mortality. Table 4 summarizes results from models where we control for the respondents' marital status. The main effect of marital status categories is highly significant and large in substantive terms: the unmarried groups experience 20% to 50% higher mortality hazard, compared to married individuals. Controlling for marital status has only a marginal impact on the main effect of education: hazard ratios for different educational levels attenuate between 0% and 3–4%, with the latter occurring at the lowest schooling levels. The education-by-female coefficients, very close to zero to begin with, also diminish only slightly. This slight attenuation, however, suffices to make the results statistically null: controlling for marital status, men and women evidence no difference in the education-mortality association. Table 4 shows results for all cohorts combined but the null findings were obtained in cohort-stratified models as well (results available from authors).

DISCUSSION

Are educational inequalities in mortality similar for U.S. men and women? Existing literature does not offer a consistent answer, although this is a fundamental issue for understanding the

causal pathways through which education affects health. This paper provides a systematic and detailed analysis of the effects of education on white and black men's and women's mortality for adults aged 25–80 and across a set of three birth cohorts.

We found that relative differences in all-cause adult mortality by education are generally comparable for men and women. This finding is largely consistent with some previous U.S. studies that also found no significant gender differences in the overall education-mortality association (McDonough et al., 1999; Zajacova, 2006). Indeed, black adults evidenced no gender differences in the education-mortality association in any birth cohort. Among white adults, some gender differences in the education-mortality association were observed among older birth cohorts. White men evidenced a tendency toward stronger education gradients at and above the level of a Bachelor's degree (16 years of schooling and above), compared to women. Although the differences were statistically significant only among the older cohorts, the point estimates for all cohorts were all in the same direction, suggesting that the mortality reduction for men at the highest levels of education is greater than that for women. A similar pattern was observed by Christenson and Johnson (1995) using Michigan data with trichotomized education: men benefited more from postsecondary education than women in terms of reduced mortality risks.

What might account for the general similarity in the U.S. education-mortality relationship between men and women across most of the educational distribution? The first part of the explanation may be the comparable content and quality of education for U.S. women and men. Holding age and other sociodemographic factors constant, boys and girls generally attend the same classes in the same schools, especially prior to postsecondary education where the academic tracks become more gendered. However, even if the human capital accumulated by men and women through education is comparable, the causal pathways that link education to health and mortality outcomes are extremely complex and may differ by gender. They include economic factors like employment, income and wealth trajectories, psychosocial factors like self-efficacy and social support, and health behaviors (Mirowsky & Ross, 2003; Ross & Wu, 1995). It is unlikely that all these pathways exhibit similar returns to education for men and women. We suspect that the observed gender similarity in the education-mortality relationship may occur because some pathways show stronger returns for men, other for women, and their effects balance out. Untangling such potentially differential mechanisms is an important goal for future research in this area.

How do we reconcile the similarity between U.S. men and women in the education-mortality association with the conclusions from European literature (i.e., Mackenbach et al., 1999) that men's education gradients in mortality are steeper than women's? First, the measurement of education differs across studies and across countries, so even a given number of years of education may not carry the same meaning in the United States as it does in various European countries. Indeed, it is a key strength of our study that we were able to separate groups with 16 years and 17–20 years of education—and this is where some gender differentials turned up. Second, the causal links between education and health likely differ across countries. For instance, labor market returns to education for men and women show large variation across context (Psacharopoulos & Patrinos, 2002) and other pathways such as health care and health behaviors vary as well. Moreover, the structure of causes of death differs across countries (Murray & Lopez, 1997). Since education is known to have different associations with mortality by cause of death (Regidor, Calle, Navarro, & Dominiquez, 2003), this would add to the international variation in the gender comparison of the education-mortality gradient.

Social stratification and health researchers may be particularly interested in the gender differences in mortality we uncovered at the highest educational attainment levels among white adults. One explanation of the stronger gradient for men focuses on selection. Individuals with

the highest levels of education, by definition, comprise a highly selected group. The selection, particularly pronounced for women in the oldest cohorts who completed at least a Bachelor's degree, may systematically impact their risks of dying during the follow-up period. Another explanation is causal and focuses on stronger returns at the highest levels of education for men than women. White women with the highest education levels may not have attained the occupational and social statuses commensurate with their education. This argument, however, is not consistent with findings on earnings returns to education. Econometric studies consistently find that the effect of schooling on income is *stronger* for women than for men across most age groups and degree levels (Dougherty, 2005; Hecker, 1998). Recently, DiPrete and Buchmann (2006) documented stronger returns to a college degree relative to a high school diploma for women compared to men consistently from the 1960s to the present.

Perhaps white men benefitted from their high educational levels more than women not though income but a higher propensity to marry (Qian & Preston, 1993). Marital status has a strong positive effect on health (Goldman, Korenman, & Weinstein, 1995), so men may have benefitted by being more likely to be married than women. Indeed, we found that among married adults, there was no gender difference in the education-mortality association. Additional factors—such as health behavior—also likely play a role in the gender differences. For instance, prevalence of smoking has been more closely tied to educational attainment among men than among women (Preston & Taubman, 1994). A recent examination of patterns in smoking-related causes of death among U.S. white adults showed that smoking patterns help explain education differentials in mortality for men versus women (Montez, Hayward, Brown, & Hummer, 2009).

Our study has several limitations. First, we cannot distinguish whether the observed differences among the three birth cohort groups are caused by actual cohort effects, age effects, or a combination of the two. The weaker effects of schooling among the oldest cohorts is consistent both with the age-as-leveler explanation whereby the effect of education weakens among the old (House, Lepkowski, Kinney, Mero, Kessler, & Herzog, 1994; Zajacova, Goldman, & Rodríguez, 2009), as well as with the cohort explanation whereby schooling is becoming more important for younger generations (Lauderdale, 2001; Lynch, 2003). This difficulty is not specific to our study – demographers have long discussed the age-cohort-period conundrum (Hobcraft, Menken, & Preston, 1982; Yang & Land, 2008). If education is becoming a more dominant predictor of mortality in recent cohorts, labor market trends are likely a major factor. Among younger birth cohorts, a higher level of educational attainment is increasingly more crucial for employment with high wages and low environmental risks (Juhn, Murphy, & Pierce, 1993; Lynch, 2006). Given the growing inequalities in income, as well as in access to health care, recent birth cohorts may continue to experience a further steepening of the mortality gradient by education.

Second, using years of education does not inform the potential effects of credentials (Ross & Mirowsky, 1999) that may matter beyond the number of completed years. For instance, there is a large difference in U.S. labor market returns to education between individuals who completed high school and those who earned a GED, although both groups might report 12 years of schooling (Heckman & Rubinstein, 2001). Third, while our data are ideally suited to describe education-mortality patterns, they include few of the mediators that likely play a role in their association, such as individual income, wealth, occupation, smoking and other health behaviors, or psychosocial factors such as social support. Another limitation is that even with the large size of the current data set, we were not able to study race/ethnic groups other than blacks and whites. Given rapid Hispanic population growth in the United States, future work with newer data should also analyze these relationships among Hispanic adults as well given that the education-health relationship has been shown to be weaker among U.S. Hispanics compared to non-Hispanic whites (Goldman, Kimbro, Turra, & Pebley 2006).

This study used one of the largest nationally representative data sources with mortality information available in the United States. In conjunction with previous research on gender differences in education differentials in mortality, this paper helps establish the general similarity between men and women in the effect of schooling on the relative risk of mortality across most of the educational distribution, with stronger educational effects for men than women at the highest education levels. This particular difference may be of substantive interest in future research because the overall education levels in the population are increasing. Thus a higher proportion of adults will fall into the schooling categories where, at present, men benefit more than women in the reduced mortality hazard. The next step in this line of inquiry is to begin isolating specific pathways that link education and mortality risk, to help disentangle the impact of the multiple factors that link higher levels of education to lower mortality risk for different population subgroups.

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Figure 1.

Effect of education on mortality hazard for men and women, by race and birth cohort: U.S. adults born 1906–1965

The figure shows log mortality hazard for each year of education relative to 12 years of schooling for men and women, by race and birth cohort. The error bars indicate 95% CI for each point estimate. Where the estimate or the CI is outside of the range of the y axis, it is not displayed. The models control for age, region of residence, and year of interview.

Table 1

Characteristics of the Sample at Interview and During Followup for Men and Women.

	Men	Women	Gender difference
<hr/>			
Non-Hispanic white (N=619,320)			
<hr/>			
Gender	48.2%	51.8%	N/A
Age at interview (mean, s.d.)	48.4 (14.2)	49.6 (14.8)	***
Married	79.5%	70.6%	***
Education, continuous (mean, s.d.)	13.2 (3.0)	12.9 (2.6)	***
Education, categorized			***
0–8	6.8%	6.0%	
9–11	9.2%	9.6%	
12	35.2%	41.1%	
13–15	20.9%	22.3%	
16	15.5%	12.6%	
17–20	12.5%	8.4%	
<i>Followup information</i>			
Years of followup (mean, s.d.)	9.6 (4.3)	9.7 (4.3)	***
Proportion died	12.3%	9.8%	***
Age at death (mean, s.d.)	69.6 (12.6)	72.4 (12.3)	***
Non-Hispanic black (N=104,952)			
<hr/>			
Gender	44.4%	55.6%	N/A
Age at interview (mean, s.d.)	46.1 (13.6)	46.8 (14.0)	***
Married	63.6%	43.3%	***
Education, continuous (mean, s.d.)	11.8 (3.3)	12.0 (3.0)	***
Education, categorized			***
0–8	13.5%	10.9%	
9–11	15.7%	17.3%	
12	37.2%	37.4%	
13–15	19.9%	21.1%	
16	8.3%	8.3%	
17–20	5.4%	5.0%	
<i>Followup information</i>			
Years of followup (mean, s.d.)	9.4 (4.3)	9.6 (4.3)	***
Proportion died	14.0%	10.3%	***
Age at death (mean, s.d.)	64.5 (14.1)	66.7 (14.0)	***

P<.001, two-tailed

Note: The table shows weighted means and proportions, and design-adjusted Wald and Chi square tests for gender differences in the distribution of key variables.

Table 2
All-Cause Mortality Hazard on Linearly Specified Education by Race and Birth Cohort.

	All cohorts			1906–25			1926–45			1946–65		
	HR	95% CI	HR	95% CI	HR	95% CI	HR	95% CI	HR	95% CI	HR	95% CI
Non-Hispanic white												
Education	0.94**	0.94–0.94	0.96**	0.96–0.96	0.92**	0.92–0.93	0.89**	0.88–0.89				
Education * female	1.01**	1.00–1.01	1.01*	1.00–1.01	1.00	0.99–1.01	0.99	0.97–1.01				
Female	0.62**	0.60–0.64	0.62**	0.61–0.63	0.64**	0.62–0.65	0.60**	0.57–0.63				
Age	1.09**	1.09–1.09	1.09**	1.09–1.09	1.09**	1.09–1.10	1.08**	1.08–1.09				
Age * female	1.00	1.00–1.00	1.00	1.00–1.01	1.00*	0.99–1.00	1.01	1.00–1.01				
Midwest	1.01	0.98–1.04	1.05**	1.02–1.08	0.97	0.93–1.01	0.94	0.86–1.02				
South	1.06**	1.04–1.09	1.04*	1.01–1.08	1.09**	1.05–1.13	1.08*	1.00–1.17				
West	1.01	0.98–1.05	1.02	0.98–1.06	0.97	0.92–1.02	1.07	0.98–1.17				
Year of interview	1.00	1.00–1.00	1.00*	0.99–1.00	1.00	0.99–1.00	1.00	0.99–1.01				
N of observations	619,320		99,506		201,788		318,026					
N of deaths	73,755		42,834		23,842		7,079					
Non-Hispanic black												
Education	0.96**	0.95–0.97	0.98**	0.97–0.99	0.95**	0.94–0.96	0.91**	0.90–0.93				
Education * female	1.00	1.00–1.01	1.00	0.99–1.02	1.00	0.99–1.02	0.98	0.95–1.01				
Female	0.64**	0.61–0.67	0.63**	0.59–0.68	0.68**	0.64–0.72	0.61**	0.55–0.67				
Age	1.07**	1.07–1.07	1.07**	1.06–1.08	1.06**	1.05–1.07	1.07**	1.06–1.09				
Age * female	1.00	1.00–1.00	1.00	0.99–1.01	1.01	1.00–1.02	1.01	1.00–1.02				
Midwest	1.06	0.99–1.14	1.07	0.96–1.19	1.08	0.97–1.21	1.00	0.87–1.16				
South	1.06	1.00–1.13	1.06	0.97–1.16	1.17**	1.06–1.29	0.91	0.81–1.03				
West	0.97	0.88–1.07	1.02	0.88–1.17	0.98	0.83–1.16	0.90	0.76–1.08				
Year of interview	1.00	1.00–1.01	1.00	0.99–1.02	1.01*	1.00–1.03	0.98*	0.97–1.00				
N of observations	104,394		13,394		33,314		58,244					
N of deaths	14,476		6,317		5,684		2,475					

**
P<0.01

* $P < 0.05$, 2-tailed

Models are adjusted for sampling design.

Table 3
Gender Differences in the Association Between All-Cause Mortality Hazard and Education, by Race and Birth Cohort.

All cohorts		1906–25		1926–45		1946–65	
HR	95% CI	HR	95% CI	HR	95% CI	HR	95% CI
Non-Hispanic white							
Education, 12 years as reference category							
0–8	1.29**	1.19**	1.14–1.24	1.52**	1.44–1.61	1.73**	1.47–2.04
9–11	1.24**	1.15**	1.10–1.20	1.32**	1.25–1.39	1.59**	1.42–1.77
13–15	0.92**	0.91**	0.86–0.96	0.96	0.91–1.02	0.86**	0.79–0.94
16	0.72**	0.81**	0.76–0.86	0.70**	0.66–0.75	0.55**	0.49–0.61
17–20	0.63**	0.74**	0.69–0.79	0.58**	0.54–0.62	0.49**	0.43–0.55
Education by female interaction							
0–8 * female	0.98	0.99	0.93–1.04	1.04	0.95–1.14	1.12	0.87–1.46
9–11 * female	1.03	1.00	0.94–1.07	1.13**	1.04–1.22	1.11	0.93–1.31
13–15 * female	0.98	1.01	0.95–1.09	0.95	0.88–1.03	0.95	0.83–1.10
16 * female	1.07*	1.01	0.92–1.11	1.12*	1.00–1.25	1.08	0.91–1.29
17–20 * female	1.12**	1.07	0.97–1.19	1.16*	1.03–1.32	1.16	0.96–1.41
Control variables							
Female	0.61**	0.61**	0.59–0.63	0.61**	0.59–0.64	0.58**	0.54–0.63
Age	1.09**	1.09**	1.09–1.09	1.09**	1.09–1.10	1.08**	1.08–1.09
Age by female	1.00	1.00	1.00–1.01	0.99*	0.99–1.00	1.01	1.00–1.01
Midwest	1.01	1.05**	1.01–1.08	0.96	0.92–1.00	0.93	0.86–1.02
South	1.08**	1.05**	1.02–1.09	1.09**	1.05–1.13	1.08	1.00–1.16
West	1.02	1.03	0.99–1.07	0.97	0.92–1.02	1.06	0.97–1.16
Year of interview	1.00	1.00*	0.99–1.00	1.00	0.99–1.00	1.00	0.99–1.01
Non-Hispanic black							
Education, 12 years as reference category							
0–8	1.26**	1.20**	1.07–1.34	1.34**	1.20–1.50	1.37*	1.06–1.76
9–11	1.23**	1.14	0.98–1.33	1.26**	1.13–1.40	1.32**	1.13–1.55

All cohorts			1906–25		1926–45		1946–65	
	HR	95% CI	HR	95% CI	HR	95% CI	HR	95% CI
13–15	0.87**	0.79–0.97	1.00	0.84–1.21	0.86	0.73–1.00	0.83*	0.70–0.98
16	0.69**	0.60–0.80	0.83	0.62–1.10	0.78*	0.62–0.99	0.55**	0.43–0.70
17–20	0.63**	0.54–0.74	0.88	0.66–1.16	0.67**	0.52–0.85	0.38**	0.25–0.57
Education by female interaction								
0–8 * female	0.96	0.87–1.06	1.01	0.86–1.18	0.85	0.72–1.01	1.34	0.95–1.91
9–11 * female	1.00	0.90–1.12	0.98	0.80–1.19	1.00	0.86–1.17	1.08	0.84–1.39
13–15 * female	1.03	0.89–1.19	0.90	0.70–1.16	1.02	0.80–1.28	1.15	0.91–1.46
16 * female	1.04	0.85–1.26	1.11	0.76–1.61	0.91	0.64–1.30	1.08	0.77–1.53
17–20 * female	1.02	0.79–1.30	1.08	0.71–1.63	0.87	0.61–1.24	1.03	0.52–2.05
Control variables								
Female	0.64**	0.59–0.69	0.63**	0.55–0.71	0.70**	0.63–0.78	0.57**	0.49–0.66
Age	1.07**	1.07–1.07	1.07**	1.06–1.08	1.06**	1.05–1.07	1.08**	1.06–1.09
Age by female	1.00	1.00–1.00	1.00	0.99–1.01	1.01	1.00–1.02	1.01	0.99–1.02
Midwest	1.05	1.25–1.33	1.07	0.95–1.19	1.07	0.96–1.20	0.99	0.86–1.15
South	1.08*	1.20–1.28	1.07	0.97–1.17	1.18**	1.07–1.31	0.91	0.80–1.03
West	0.99	0.89–0.95	1.02	0.89–1.18	1.01	0.85–1.19	0.91	0.76–1.09
Year of interview	1.00	0.69–0.75	1.00	0.99–1.02	1.01*	1.00–1.03	0.98*	0.97–1.00

^{**} P<0.01

^{*} P<0.05, 2-tailed

Models are adjusted for sampling design.

Table 4
Gender Differences in the Association Between All-Cause Mortality Hazard and Education Adjusting for Marital Status

	Non-Hispanic White		Non-Hispanic Black	
	OR	95% CI	OR	95% CI
Education, 12 years as reference category				
0–8	1.26**	1.22–1.30	1.22**	1.13–1.31
9–11	1.23**	1.19–1.27	1.21**	1.11–1.31
13–15	0.92**	0.89–0.95	0.88*	0.79–0.97
16	0.72**	0.69–0.75	0.70**	0.60–0.81
17–20	0.63**	0.60–0.65	0.64**	0.55–0.76
Education by female interaction				
0–8 * female	0.98	0.93–1.02	0.96	0.87–1.07
9–11 * female	1.02	0.98–1.08	1.00	0.90–1.12
13–15 * female	0.98	0.93–1.02	1.02	0.88–1.19
16 * female	1.06	1.00–1.13	1.04	0.85–1.27
17–20 * female	1.07	1.00–1.16	1.00	0.78–1.29
Marital status (ref.=married)				
Divorced	1.49**	1.44–1.53	1.27**	1.20–1.35
Widowed	1.25**	1.22–1.27	1.20**	1.13–1.28
Never married	1.49**	1.44–1.55	1.55**	1.45–1.65
Unknown marital status	0.97	0.75–1.26	0.96	0.57–1.61
Control variables				
Female	0.60**	0.58–0.62	0.60**	0.56–0.65
Age	1.09**	1.09–1.09	1.07**	1.07–1.07
Age by female	1.00**	1.00–1.00	1.00	1.00–1.00
Midwest	1.01	0.99–1.04	1.06	0.99–1.14
South	1.09**	1.06–1.12	1.10**	1.03–1.17
West	1.02	0.99–1.05	1.00	0.91–1.11
Year of interview	1.00	1.00–1.00	1.00	1.00–1.01

**
P<0.01

*
P<0.05, 2-sided

Models are adjusted for sampling design.