Gender Differences in Education Gradients in Mortality by Cause

Across 60 U.S. Birth Cohort: NHIS 1986-2002.

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Abstract.

We present detailed analyses of gender differences in the effects of education on cause-specific mortality among non-Hispanic white U.S. adults born between 1906 and 1965. The existence of education gradients in mortality has been well established. The question whether men and women differ in how education affects their mortality, however, has been neglected despite its importance for the examination of causal pathways that link schooling and death. Data from the 1986-2000 National Health Interview Surveys matched to the National Death Index through 2002 were analyzed using cohort-stratified proportional hazard models to estimate gender differences in the effect of education on the mortality hazard. Results indicate that there are few systematic differences in education inequalities in all-cause mortality except some tendency for steeper education gradients at postsecondary levels of schooling for men, compared to women. The cause-specific analyses suggest a slightly stronger gradient for women in cardiovascular and diabetes mortality; for men in respiratory mortality. For the most part, the differences are not statistically significant. Some explanations for the observed patterns are discussed.

The goal of this paper is to provide a detailed analysis of gender differences in the association between education and cause-specific mortality for non-Hispanic white U.S. men and women born between 1906 and 1965.

The strong effects of education on mortality have been described exhaustively (i.e., Pappas, Queen et al. 1993; Preston and Elo 1995; Lleras Muney 2005). It has long been known that the schooling effects may vary by population subgroups, whether by age, race, or cohort (Kitagawa and Hauser 1973; Bassuk, Berkman et al. 2002; Lynch 2006). However, we still lack evidence-supported consensus regarding how educational attainment relates to U.S. men's and women's odds of dying from major causes of death. In this paper, we focus on three empirical questions. First, do education gradients in all-cause mortality differ for men and women? Second, do the gender patterns vary by birth cohorts? And third, are the gender patterns similar across specific causes of death?

Few researchers focused on analyzing gender differences in educational gradients in all-cause mortality, and even fewer studied cause-specific mortality (Regidor, Calle et al. 2003). In many studies, however, analyses of the education-mortality association were conducted separately for men and for women and thus allow some gender comparisons. Most findings from European countries report larger education differentials in mortality for men than for women. Using death registry data from several European countries, larger relative and absolute education differentials in all-cause mortality were found for men than for women (Koskinen and Martelin 1994; Mustard and Etches 2003; Elo, Martikainen et al. 2006), although the gender differences diminished at older ages (Mackenbach, Kunst et al. 1999; Huisman, Kunst et al. 2004; Huisman, Kunst et al. 2005).

A considerable cross-country variation existed in both relative and absolute education differentials for both men and women (for example, see Huisman, Kunst et al. 2004), suggesting large differences in how a given educational level translates to mortality advantage or disadvantage across national contexts. In particular, the education gradient is steeper in the U.S. than in most European countries (Kunst and Mackenbach 1994). Hence, we cannot assume that the results from Europe will automatically apply to the U.S. context because the returns to education in earnings and other factors differs cross-nationally, and even the meaning of a given number of schooling years may vary intenrationally (Card 1999).

In the U.S., the patterns observed for men's and women's education gradients in all-cause mortality are less consistent. A number of studies found larger education differentials in all-cause mortality for men than for women (Feldman, Makuc et al. 1989; Pappas, Queen et al. 1993; Mackenbach, Kunst et al. 1999; Bassuk, Berkman et al. 2002; Molla, Madans et al. 2004). Other researchers reported little or no gender differentials (Feldman, Makuc et al. 1989; Sorlie, Backlund et al. 1995; Elo, Martikainen et al. 2006). Finally, some analyses found slightly larger relative educational differences in mortality for women than men (Deaton and Paxson 1999), including the seminal work by Kitagawa and Hauser (1973).

It is important to note that the above studies did not test whether the observed gender differences were statistically significant. We found only three studies that explicitly tested the significance of the gender difference in the effect of education on all-cause mortality. All used U.S. data. Christenson and Johnson (1995) reported substantively small but statistically significant difference in the education gradients for men and women. The differences were in the opposite direction at different education levels: men evidenced a stronger gradient at the postsecondary levels, women at the pre-secondary levels. Two additional studies used nationally representative data with a linear specification of education. Both reported no significant gender difference in the effect of education on all-cause mortality (McDonough, Williams et al. 1999; Zajacova 2006).

In order to move research toward understanding the causal effects of schooling, we need to study specific causes of death. Not only does the distribution of the causes differs by sex (Case and Paxson 2005), but also the different causes appear to have varying associations with education. For instance, breast cancer has a direct education gradient, in contrast to most other causes (Steenland, Henley et al. 2002; Regidor, Calle et al. 2003). Similarities or differences between men and women in education gradients for specific causes of death may provide important insights into the specific mechanisms that link schooling and health.

In the published studies, some causes of death have a stronger gradient for men, others for women, and some seem to show no gender differences in the education effects. Cardiovascular disease mortality is the leading cause of death, and it is also the one cause for which there appears to be a relatively consistent gender pattern. In a number of European and U.S. studies, women across most or all age groups were found to have larger education differentials compared to men (Mackenbach, Kunst et al. 1999; Steenland, Henley et al. 2002; Regidor, Calle et al. 2003; Huisman, Kunst et al. 2005; Elo, Martikainen et al. 2006). Another cause of death with possibly stronger education gradients for women is diabetes. Several studies found diabetes mortality is strongly dependent on education for women, while the gradient was relatively flat for men (Steenland, Henley et al. 2002; Regidor, Calle et al. 2003).

The gender patterns appear reversed for cancer mortality. Most researchers find that cancer mortality has a stronger education gradient for men than women (i.e. Elo, Martikainen et al. 2006; Albano, Ward et al. 2007). This appears to be in part due to lung cancer, which has a very strong association with education for men and less so for women (Mackenbach, Kunst et al. 1999; Huisman, Kunst et al. 2005; Albano, Ward et al. 2007). Other smoking-related respiratory diseases, such as COPD, also have a stronger education gradient for men than for women (Mackenbach, Kunst et al. 1999; Steenland, Henley et al. 2002; Huisman, Kunst et al. 2005). External deaths comprise another major category where education has been show more strongly related to mortality for men at most or all ages, compared to women (Mackenbach, Kunst et al. 1999; Huisman, Kunst et al. 2005; Elo, Martikainen et al. 2006).

Finally, no gender differences are evident for cerebrovascular mortality across different age groups (Avendaño, Kunst et al. 2004; Huisman, Kunst et al. 2005), although Mackenbach et al.'s (1999) results suggest slightly larger differentials for men.

The cursory review above described the current state of developments in this field of research. This paper extends the work by a systematic survey of education gradients in cause-specific mortality using a large, nationally representative sample of U.S. adults. We formally test for gender differences in education gradients in all-cause and cause-specific mortality across a wide range of birth cohorts.

DATA AND METHODS

Data

The analyses are based on the National Health Interview Survey (NHIS) data for years 1986-2000, linked to multiple cause-of-death files from the National Death Index that includes deaths through 2002. 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 rate*s* were over 95% in the earlier years (Massey, Moore et al. 1989) and declined to 89% in 2000 (CDC 2002).

Additional information about the sampling design are available in Massey et al. (1989) for years 1985-1994 and in Botman et al. (2000) for 1995-2004. The completed data and accompanying documentation for all survey years are available on the NCHS website: http://www.cdc.gov/nchs/about/major/nhis/quest_data_related_doc.htm.

Vital status for adults who participated in 1986-2000 NHIS was determined by a probabilistic match to the National Death Index (Lochner, Hummer et al. 2007).

Sample

We include non-Hispanic white respondents who were born between 1906 and 1965 and were 25-80 years at the time of the interview. The lower age boundary was selected as the age where most individuals have completed their schooling. The upper age boundary was set because of indications that the matching of death information was less successful for women over 80 compared to men (Masters, Brown and Hayward, unpublished analyses), which could lead to biased findings. Individuals who self-reported as white and as 'not Hispanic' were coded as non-Hispanic whites. There were no missing values on age, sex, and region of residence. Individuals with missing values on education (0.6 percent) were not included in the analyses. Because of the very low proportion of missing observations in the variables used in the analyses, we do not believe the missingness has the potential to bias our findings.

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 12^{th} year and in achieved postsecondary degrees. We converted the degrees into years of schooling as follows: 'some college, 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 the analysis. In cause-specific analyses we used education as a trichotomized predictor: less than high school (<12 years), high school as reference (12 years), more than high school (>12 years). In all-cause mortality we estimated models with education as dummies for each year of schooling (reference category=12 years), and as a linear predictor centered on 12 years. The mortality data file included the underlying cause of death and select contributing causes, coded using the ICD-10 classification. We categorized causes into one of the following major categories: cardiovascular, cerebrovascular, diabetes, cancer excluding lung, lower-respiratory disease including lung cancer, external, and other causes. Lung cancer was included among lower-respiratory diseases such as COPD or emphysema because these causes are all strongly affected by smoking behaviors.
Control variables included age centered on a model-specific mean, region of residence (Northeast as reference), and year of interview (centered on the mean year 1993).

Analysis.

We estimated proportional hazard models to calculate mortality hazard ratios by educational status for men and women. We employed gender-stratified models and interaction models with an education-by-gender coefficient. The statistical significance of the interaction coefficient was used to formally test for gender differences in the effects of education on the mortality hazard. It is known that the effect of some control variables, especially age, on mortality differs by gender (Malyutina, Bobak et al. 2004; Arias 2006). In order to approximate the flexibility of gender-stratified models to capture the data structure for each gender, the interaction models included an age by female interaction.

Most analyses were stratified by birth cohort: 1906-1925, 1926-1945, and 1946-1965. All models adjusted for age at interview, year of interview, gender, and region of residence. Continuous predictors were included in the model centered around their approximate means. All analyses adjusted for sampling design by employing the suite of survey commands in Stata 10 (StataCorp 2007).

RESULTS

Sample description.

Table 1 shows the characteristics of the sample at interview and mortality status information during followup. There was about an equal number of men and women. Their mean age at interview was 49 years and their mean year of birth was 1944. About 46% had some postsecondary education, 38% completed 12 years, and about 16% did not complete high school. About 11% died during the followup period that averaged 10 years. The mean age of death was 71 years. The modal cause of death was cardiovascular disease and two other major causes are cancer and respiratory disease for both men and women.

All-cause mortality.

The first step is to determine the shape of the education-mortality association, using proportional hazard models with dummies for single years of education (12 as reference). The results are used to select the most appropriate functional form for the education variable in subsequent analyses. Figure 1 shows the effect of education on all-cause mortality by gender and cohort, using single years of schooling as well as a linear specification of education. The figure highlights three issues: 1) The effects of education become unreliably estimated at the lowest levels of education, due to the relatively sparse data at those attainment levels; 2) the effect of education does not include clear nonlinearities , whether marked drops in the mortality hazard at the credentialing years or a substantially different slope for education at different levels. This indicates that a linear specification of education would not be improper. However, a number of previous studies suggested that the education effect for men vs. for women may differ at the pre-secondary vs. postsecondary levels (i.e., Christenson and Johnson 1995). The categorized specification may also be more informative for hypothesis generation for future studies concerning the mediators of the education-mortality association. 3) The effect of education is weaker for older cohorts and strengthens for younger cohorts. Within each cohort, however, the effect of education for men and women looks very similar, in both the dummy and linear specifications.

To formally test for differences in the education-mortality association, we present results from Cox models of all-cause mortality that include the interaction term of education with gender. The results are shown in table 2 for linear education and table 3 for categorized education. The first, linear, specification was selected for parsimony. The 6-category education models allow testing whether any gender differences exist at specific educational levels.

The main effect of education is strong: for the sample as a whole (column 1), each year of education is associated with about 6% lower mortality hazard for men. There is some evidence of a weaker effect of education for women although the difference is substantively small and appears to be driven by the patterns in the oldest birth cohort. The two younger birth cohorts evidence no significant gender difference in the effect of schooling – if anything, the coefficient is in the direction of a stronger effect for women relative to men.

To evaluate gender difference separately at various educational attainment levels, we present in table 3 the all-cause mortality models with a 6-category education. The results show a clear cohort trend whereby the education-mortality association is much stronger for the youngest cohorts, compared to their older counterparts. The main effect of education is strong – with one exception (some college compared to the baseline high school for the cohort born 1926-45), all coefficients are highly significant. However, there is relatively little difference between men and women. In the first column with all birth cohorts combined, women evidence a significantly larger difference between some college vs. high school than men, but a significantly smaller difference between a Bachelor's degree vs. high school than men. When disaggregated by birth cohort, this pattern does not hold in any of the three groups. Only one of the coefficients, comparing 9-11 years vs. high school in the middle birth cohort, is significant at $p<.01$.

Specific causes of death.

The all-cause mortality trends may be obscuring differences in specific causes of death. Table 4 shows results from interaction models of major causes of death on trichotomized education, adjusting for the full set of control variables. For parsimony, the table lists only the main coefficients of interest: the main effect on mortality hazard of less than high school and more than high school, relative to high school (for men) and the two education by gender interaction coefficients.

The results show some differences across the specific causes. In mortality from cardiovascular disease, odds of dying for women with less than high school relative to high school are substantively and statistically significantly higher than for men in all cohorts. There is some trend toward the opposite gender pattern at the post-secondary level, with stronger education gradients for men compared to women.

For respiratory diseases, there is an interesting reversal of the gender patterns across cohorts. Among the oldest cohort, women evidence a significantly flatter education gradient than men. Among the youngest cohort, in contrast, women evidence a significantly stronger education gradient compared to men, at both the pre-secondary and post-secondary levels. This trend could be explained by gender differences in the uptake of smoking across birth cohorts by socioeconomic status.

For diabetes, cerebrovascular, external, causes, there were no or few statistically significant differences between men and women in how education and mortality are associated – given the number of estimated coefficients, some of the significant results could be expected by chance alone. In most cases, the point estimates of the interaction coefficients were also small – although note the results for diabetes that suggest stronger education gradients for women than for men.

Figures 2a, b, and c illustrate these findings graphically for three major causes: cardiovascular, respiratory, and cancer. The high relative differences for women at the youngest cohorts are partially due to the low baseline mortality in that cohort, coupled with a trend toward premature mortality among the least-educated women.

We completed a number of diagnostic tests to assess the validity of our findings. To examine whether the effect of education on mortality varied with the length of followup or with age, we estimated models where education was allowed to vary with time (years and log years) and with age. The results showed that the effect of education weakened slightly with time and age, more so for the oldest cohort and remaining stable for the youngest cohorts, among both men and women. The next set of tests comprised survival models with frailty using different specification for the distribution of unobserved heterogeneity. These models suggested the presence of a non-zero distributed frailty but the coefficients for education strengthened only marginally. Our results were robust to different estimation methods and did not differ substantively whether we controlled for sampling design or not. Finally, to test the proportionality of hazards, we plotted Kaplan-Meier survival curves and log-log plots by gender, education, and cohort. There appeared to be a slight narrowing of mortality hazards with time, a finding consistent with the interaction models of education with time or age. However, in no case did we find severe departures from the proportionality assumption.

DISCUSSION

Are educational inequalities in mortality similar for U.S. men and women? Little previous research informs this question, although this is a fundamental issue in terms of our understanding of the causal pathways through which education affects health. Even less is known about specific causes of death. This paper provides a systematic, thorough analysis of the effects of schooling on men’s and women’s mortality across a wide range of birth cohorts and for specific causes of death.

Our analyses indicate that the effects of education on all-cause mortality are generally comparable for men and women, with some tendency toward stronger gradient for men at the post-secondary level, especially above a Bachelor's degree. We identified only three previous studies that tested for gender differences in the education-mortality association for U.S. adults (Christenson and Johnson 1995; McDonough, Williams et al. 1999; Zajacova 2006). Our findings partially corroborate their conclusions of little systematic gender differences in the effects of education (McDonough and Berglund 2003; Zajacova 2006), as well as the stronger education inequalities for men with more than a HS diploma (Christenson and Johnson 1995). The gender patterns look similar across 60 birth cohorts. The education gradients vary strongly by cohort, whereby older cohorts evidence much flatter gradients, compared to younger cohorts. Across the different cohorts, however, men's and women's education effects look similar, a remarkable stability given the societal changes that occurred during this time in employment, earnings, marital status, health behaviors, and other potential mediators of the education-mortality association.

Analyses by specific causes of death showed some variability. For cardiovascular mortality, women evidenced a stronger education gradient at the pre-secondary levels and a tendency toward a weaker gradient at the post-secondary levels in all cohorts. Since cardiovascular mortality is the major cause of death, it suggests that the all-cause results are driven by the patterns observed for this cause. Respiratory mortality shows a remarkable flip from flatter gradients for women in older cohorts to stronger ones in younger cohorts. This change is likely due to the patterns in smoking by women vs. men across different periods and socioeconomic classes. In other categories, the gender differences were not statistically significant.

It is important to note that, whether statistically significant or not, the point estimates that measure the gender differences in education effects for specific causes of death are not in the same direction. For instance, for cerebrovascular mortality, the data suggest a tendency toward a flatter education gradient at both pre- and post-secondary levels for women. In contrast, for cancer mortality, the data suggest a tendency toward a stronger gradient at both schooling levels for women. While by definition these tendencies may have occurred entirely by chance, they could point toward different mechanisms by which education affects health for men and for women.

Our study has several limitations. First, we cannot distinguish whether the observed cohort differences (weaker education effects for older cohorts) are due to true differences across birth cohorts in how education affects health, or whether the flatter gradients for the older individuals are an artifact of mortality selection (see Lauderdale 2001; Lynch 2003). We can only claim that the patterns we observe are conditional to survival to the age at interview and are period-specific. Second, we do not know the extent of random or systematic measurement error in classifying the causes of death. It is known that the assessment of the underlying causes of death is often imprecise, even in developed countries (Mathers, Ma Fat et al. 2005). This may be less of an issue for some causes (external, cancer) but more for others such as diabetes. And finally, while our data are ideally suited to describe the 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.

Table 1. Characteristics of the sample at interview and vital status during followup, non-Hispanic whites, NHIS 1986-2002.

	Men	Women	Total
N (by gender as proportion of total)	294,125(48.2%)	325,195(51.8%)	619,320
Age at interview	48.4	49.6	49.0
Birth cohort	1944	1943	1944
Education			
Less than high school	15.9%	15.6%	15.8%
High school completion	35.2%	41.1%	38.2%
More than high school	48.9%	43.3%	46.0%
Region of residence			
Northeast	21.6%	21.9%	21.7%
Midwest	27.3%	27.1%	27.2%
South	32.5%	32.7%	32.6%
West	18.6%	18.3%	18.5%
Followup information			
Proportion died	12.3%	9.8%	11.0%
Years of followup (mean, s.d.)	9.6	9.7	9.7
Age at death (mean, s.d.)	69.6	72.4	70.9
Cause of death (as proportion of total deaths)			
Cardiovascular	34.7%	31.3%	33.1%
Cerebrovascular	5.0%	7.3%	6.0%
Respiratory (including lung cancer)	16.1%	14.2%	15.2%
Cancer (except lung)	19.3%	22.0%	20.6%
Diabetes	2.5%	3.1%	2.8%
External	5.6%	3.3%	4.6%
Other	16.8%	18.8%	17.7%

Note: Adjusted for sampling design except for sample size information.

Table 2. All-cause mortality hazard on linearly specified education by birth cohorts, for NH white men and women.

	All cohorts	1906-25	1926-45	1946-65
Female	0.620***	0.619***	0.637***	0.600***
	(0.0080)	(0.0062)	(0.0088)	(0.017)
Age	1.089***	1.091***	1.093***	1.083***
	(0.00053)	(0.0018)	(0.0017)	(0.0031)
Age by female	1.000	1.002	0.995**	1.006
	(0.00068)	(0.0023)	(0.0021)	(0.0042)
Midwest	1.010	1.049***	0.966	0.938
	(0.014)	(0.017)	(0.020)	(0.040)
South	1.064***	1.043*	1.086***	1.083*
	(0.014)	(0.017)	(0.021)	(0.042)
West	1.014	1.022	0.973	1.066
	(0.017)	(0.020)	(0.025)	(0.048)
Year of interview	0.999	0.995*	0.996	1.002
	(0.0015)	(0.0020)	(0.0028)	(0.0049)
Education	0.942***	0.961***	0.924***	0.885***
	(0.0015)	(0.0020)	(0.0025)	(0.0046)
Education * female	1.007***	1.006*	0.997	0.989
	(0.0024)	(0.0031)	(0.0049)	(0.0096)
N				

** p<.05 ***p<.01

Note: Mortality hazard ratios and standard error shown. Adjusted for sampling design except for N, which represents the actual number of observations in the dataset.

Table 3. All-cause mortality hazard on categorized education by birth cohorts, for NH white men and women.

	All cohorts	1906-25	1926-45	1946-65
Female	0.609***	0.610****	0.613***	0.581***
	(0.010)	(0.011)	(0.013)	(0.024)
Age	1.090***	1.091***	1.093***	1.084***
	(0.00054)	(0.0018)	(0.0017)	(0.0031)
Age by female	1.000	1.002	0.995**	1.005
	(0.00068)	(0.0023)	(0.0021)	(0.0042)
Midwest	1.006	1.046***	0.963	0.934
	(0.013)	(0.017)	(0.020)	(0.040)
South	1.076***	1.055***	1.089***	1.075
	(0.014)	(0.017)	(0.021)	(0.042)
West	1.018	1.028	0.973	1.062
	(0.017)	(0.021)	(0.025)	(0.048)
Year of interview	0.999	0.995*	0.997	1.002
	(0.0015)	(0.0020)	(0.0028)	(0.0049)
Education, 12 years as reference category				
0-8	1.288***	1.187***	1.518***	1.731***
	(0.021)	(0.025)	(0.043)	(0.15)
9-11	1.241***	1.150***	1.321***	1.589***
	(0.022)	(0.027)	(0.036)	(0.088)
13-15	0.922***	0.911***	0.964	0.861***
	(0.016)	(0.025)	(0.026)	(0.037)
16	0.717***	0.812***	0.701***	0.547***
	(0.015)	(0.025)	(0.023)	(0.029)
17-20	0.627***	0.737***	0.582***	0.490***
	(0.013)	(0.024)	(0.020)	(0.030)
0-8 * female	0.982	0.986	1.040	1.123
	(0.022)	(0.028)	(0.047)	(0.15)
9-11 * female	1.030	1.004	1.129***	1.108
	(0.025)	(0.032)	(0.045)	(0.096)
13-15* female	0.983**	1.014	0.950	0.953
	(0.025)	(0.036)	(0.040)	(0.068)
16 * female	1.071***	1.012	1.122*	1.084
	(0.034)	(0.047)	(0.063)	(0.097)
17-20 * female	1.123	1.072	1.163*	1.164
	(0.043)	(0.057)	(0.074)	(0.12)
N				

** p<.05 ***p<.01

Note: Mortality hazard ratios and standard error shown. Adjusted for sampling design except for N, which represents the actual number of observations in the dataset.

Table 4. Cause-specific mortality hazard on trichotomized education by birth cohorts, for NH white men and women.

	All cohorts	1906-25	1926-45	1946-65
Cardiovascular				
Education for males, HS as reference				
Less than HS	1.266***	1.184***	1.445***	1.640***
	(0.029)	(0.037)	(0.055)	(0.16)
More than HS	0.763***	0.824***	0.735***	0.609***
	(0.019)	(0.029)	(0.029)	(0.046)
Education by gender				
Less than HS * female	1.127***	1.124***	1.193***	1.610***
	(0.037)	(0.047)	(0.075)	(0.28)
More than HS * female	1.087*	1.052	1.033	1.343*
	(0.037)	(0.048)	(0.069)	(0.19)
Cerebrovascular				
Education for males, HS as reference				
Less than HS	1.215***	1.129	1.379***	1.992**
	(0.076)	(0.085)	(0.153)	(0.593)
More than HS	0.759***	0.752***	0.790**	0.750
	(0.047)	(0.058)	(0.091)	(0.176)
Education by gender				
Less than HS * female	0.924	0.985	0.788	0.851
	(0.076)	(0.095)	(0.135)	(0.382)
More than HS * female	1.143	1.164	1.122	0.765
	(0.092)	(0.118)	(0.182)	(0.270)

Table 4 continued.

Table 4 continued

Respiratory				
Education for males, HS as reference				
Less than HS	1.453***	1.349***	1.675***	1.838***
	(0.048)	(0.061)	(0.077)	(0.35)
More than HS	0.641***	0.674***	0.664***	0.503***
	(0.025)	(0.037)	(0.037)	(0.074)
Education by gender				
Less than HS * female	0.874***	0.780***	1.063	1.712*
	(0.041)	(0.051)	(0.074)	(0.43)
More than HS * female	1.130*	1.236***	1.216*	0.878
	(0.062)	(0.092)	(0.10)	(0.20)
Cancer (except lung)				
Education for males, HS as reference				
Less than HS	1.092*	1.057	1.196***	1.239
	(0.038)	(0.046)	(0.061)	(0.15)
More than HS	0.877***	0.945	0.873***	0.715***
	(0.029)	(0.044)	(0.040)	(0.061)
Education by gender				
Less than HS * female	0.935	0.895	1.016	1.051
	(0.042)	(0.053)	(0.075)	(0.17)
More than HS * female	0.993	0.985	1.045	1.083
	(0.045)	(0.065)	(0.066)	(0.12)

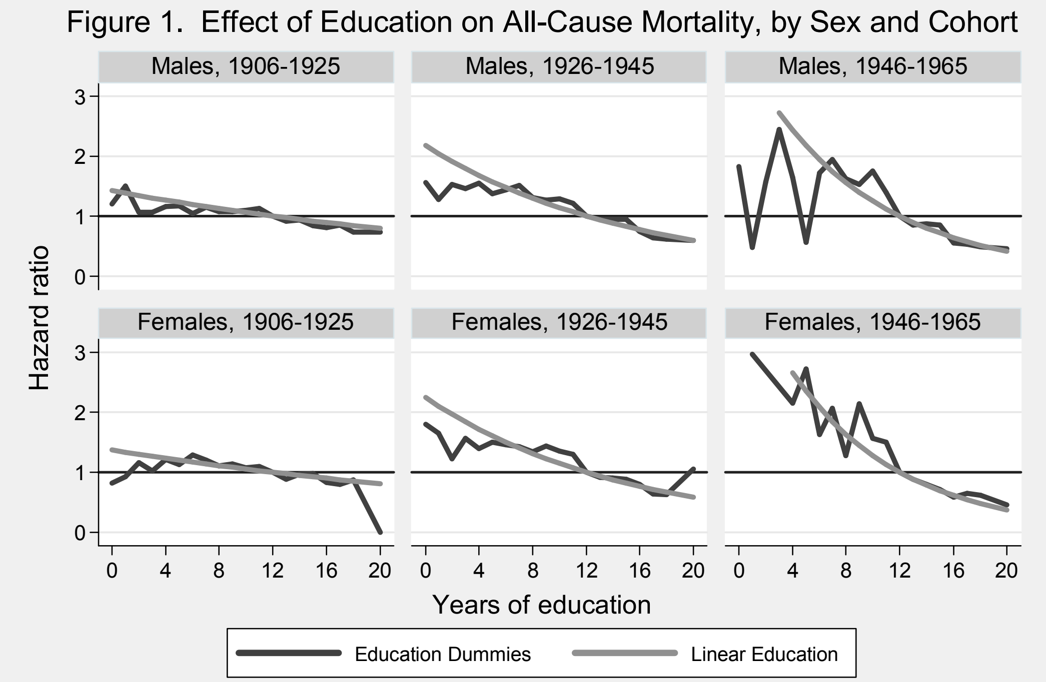
Table 4 continued

Table 4 continued

Diabetes				
Education for males,				
HS as reference				
Less than HS	1.212*	1.244	1.214	1.572
	(0.11)	(0.15)	(0.16)	(0.49)
More than HS	0.737***	0.955	0.624***	0.570*
	(0.070)	(0.14)	(0.083)	(0.14)
Education by gender				
Less than HS * female	1.265*	1.024	1.881***	0.971
	(0.15)	(0.16)	(0.35)	(0.42)
More than HS * female	0.857	0.754	1.170	0.633
	(0.12)	(0.16)	(0.24)	(0.24)
External				
Education for males,				
HS as reference				
Less than HS	1.371***	1.177	1.310*	1.587***
	(0.081)	(0.12)	(0.14)	(0.14)
More than HS	0.703**	0.790	0.773*	0.649***
	(0.037)	(0.099)	(0.080)	(0.045)
Education by gender				
Less than HS * female	1.005	1.014	1.115	0.887
	(0.10)	(0.16)	(0.21)	(0.17)
More than HS * female	1.236*	1.247	1.035	1.242
	(0.12)	(0.22)	(0.19)	(0.18)

** p<.05 ***p<.01

Note: All models control for age, age by female interaction, female, region of residence, and year of interview. Full results available from authors.



Results from models by sex, cohorts, and two different specifications of education.

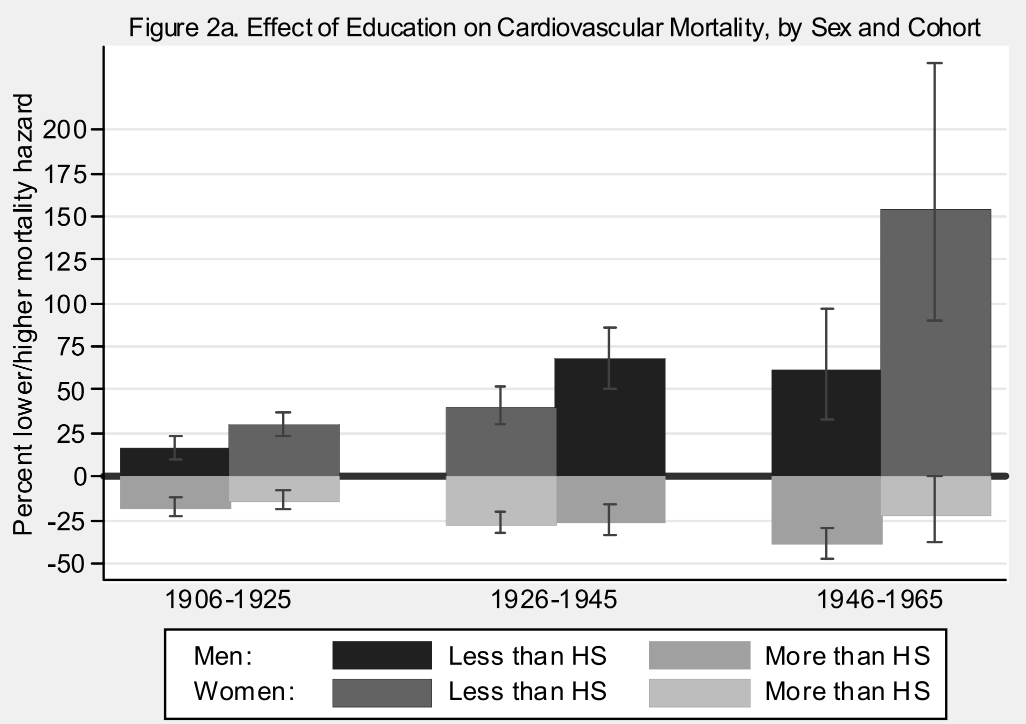
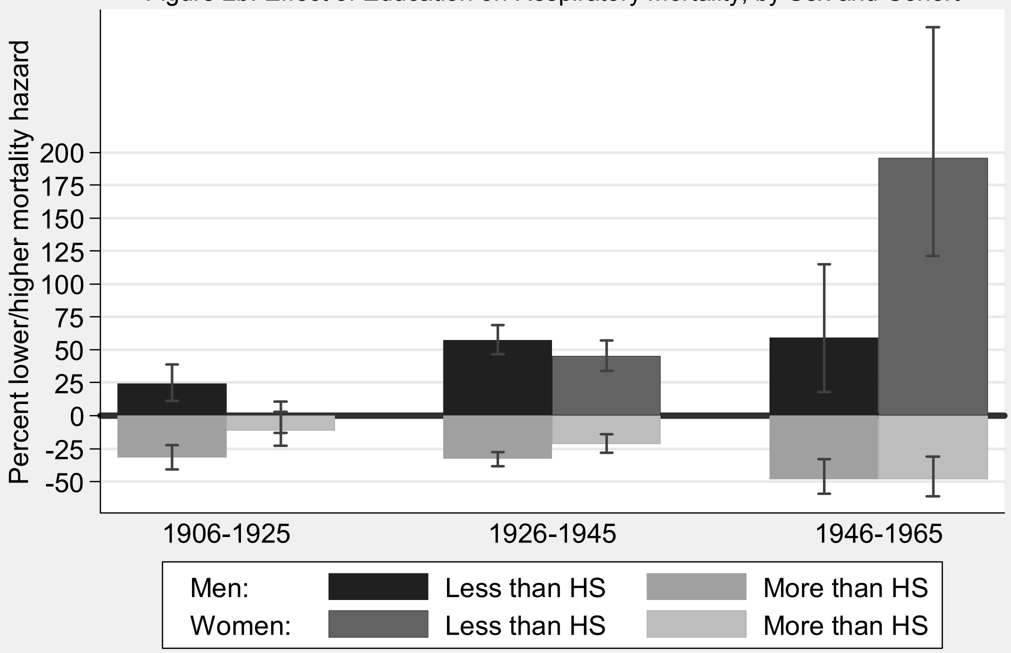
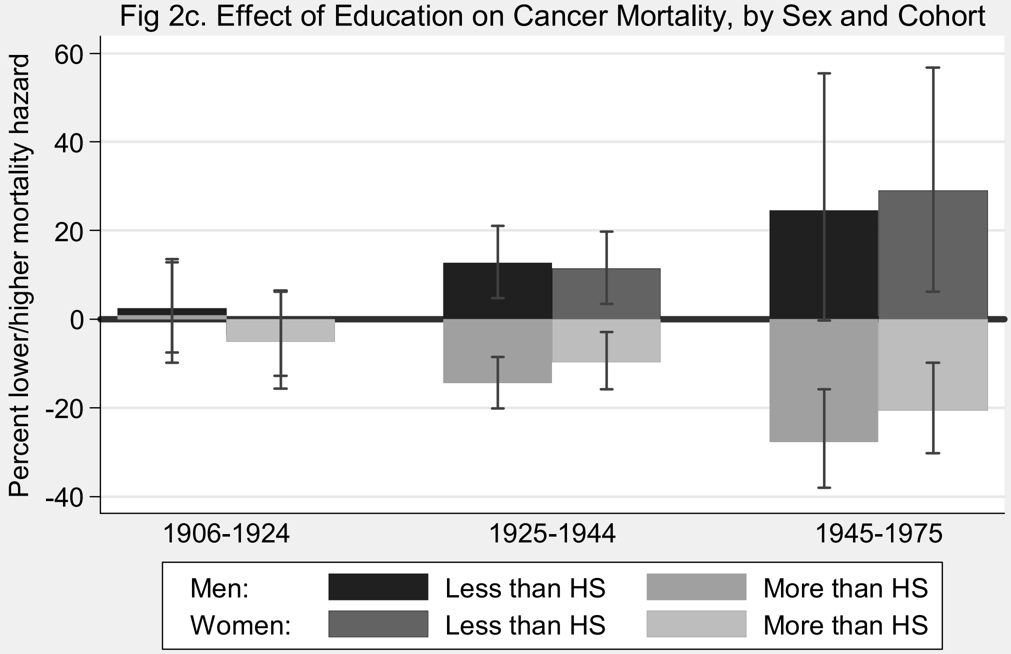


Figure 2b. Effect of Education on Respiratory Mortality, by Sex and Cohort



Results from respiratory-deaths survival models stratified by sex and cohort.

Fig 2c. Effect of Education on Cancer Mortality, by Sex and Cohort



Results from cancer-deaths survival models stratified by sex and cohort.
Excludes respiratory cancers.

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