Why Is the Educational Gradient of Mortality Steeper for Men?

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Objectives. It is often documented that the educational gradient of mortality is steeper for men than for women; yet, the explanation remains a matter of debate. We examine gender differences in the gradients within the context of marriage to determine whether overall differences reflect gender differences in health behaviors or a greater influence of men's education on spousal health.

Methods. We used data from the 1986 through 1996 National Health Interview Survey Linked Mortality Files for non-Hispanic White adults aged 55–84 years at the time of survey. We estimated Cox proportional hazards models to examine the gradients ($N = 180,208$).

Results. The educational gradient of mortality is marginally steeper for men than for women when aggregating across marital statuses; yet, this reflects a steeper gradient among unmarried men, with low-educated never married men exhibiting high levels of mortality. The gradient among unmarried men is steeper than unmarried women for causes that share smoking as a major risk factor, supporting a behavioral explanation for differences in the gradient. No gender difference in the gradient is observed for married adults.

Discussion. Low education and unmarried status exert a synergistic effect on men's mortality. Unmarried, low-educated men may lack social supports that encourage positive health behaviors.

Key Words: Education—Gender—Gradient—Marriage—Mortality.

THE inverse association between educational attainment and adult mortality in the United States is a social fact. Further, prior studies frequently report that the educational gradient appears steeper for men than for women (Elo & Preston, 1996; Feldman, Makuc, Kleinman, & Cornoni-Huntley, 1989; Jemal et al., 2008; Lin, Rogot, Johnson, Sorlie, & Arias, 2003; Molla, Madans, & Wagener, 2004; Nathanson & Lopez, 1987; Preston & Taubman, 1994; Rogot, Sorlie, & Johnson, 1992; Singh & Siahpush, 2001). However, statistical tests of this visually impressive difference are rarely conducted. Among four studies that statistically tested for gender differences in the educational gradient of mortality in the United States, two found no difference using a linear measure of education (McDonough, Williams, House, & Duncan, 1999; Zajacova, 2006), whereas two reported a marginally steeper gradient for men using a categorical measure of education, with mortality reduction larger for men at the postsecondary level (Christenson & Johnson, 1995; Zajacova & Hummer, under review).

A gender difference in the gradient potentially signals differential returns to education due to a host of factors, including biomedical, behavioral, and/or social structural differences in the lives of men and women (Nathanson & Lopez, 1987). Not surprisingly, given the absence of in-depth attention to this issue, substantial ambiguity surrounds which of these factors come into play. In this paper, we build on the work of Nathanson and Lopez (1987) and Smith and Waitzman (1994) by evaluating gender differences in the educational gradient of mortality within the context of marriage. We disaggregate the gradients by marital status because education and marital status are fundamental determinants of life chances, and the association each has with mortality may depend on the presence of the other. We begin by formally testing for gender differences in the gradient between non-Hispanic White men and women 55 years of age and older. Where differences exist, we examine leading causes of death to evaluate the role of health behaviors that may have contributed to those differences. We also evaluate the possibility that the steeper gradient among men reflects a disproportionate influence of their education on household resources and health.

Background

At least as early as 1960, gender differences in the educational gradients were apparent, although at that time the gradient appeared steeper for women than for men across most age groups (Kitagawa & Hauser, 1973). Subsequent studies on temporal changes in the gradient between 1960 and the mid-1980s revealed its malleability. By the mid-1980s, the gradient had become much steeper among men (Crimmins & Saito, 2001; Lauderdale, 2001; Pappas, Queen, Hadden, & Fisher, 1993; Preston & Elo, 1995; Rogot et al., 1992), largely due to disproportionate declines in heart disease mortality among higher-educated men (Feldman et al., 1989). Yet, among women it is unclear

whether their gradient became steeper (Lauderdale, 2001; Pappas et al., 1993), flatter (Rogot et al., 1992), was unchanged (Feldman et al., 1989) or changed in a cohort-specific fashion (Crimmins & Saito, 2001; Preston & Elo, 1995). The net result of these gender-specific trends during this period was a visually steeper educational gradient among men in comparison to women across adulthood.

The seminal article on this issue by Nathanson and Lopez (1987) introduced the idea that these gender differences in the gradient likely have a behavioral explanation that may be revealed if the gradients are assessed within the context of marriage. Drawing on 1971 data from Canada, they showed that the gradient for mortality—indexed by income instead of education—was steeper for men than for women largely due to high mortality rates among low-income men. Nathanson and Lopez speculated that compared with women with low income or other men, these men may be less likely to have access to health-enhancing social ties such as marriage and, therefore, are more likely to engage in risky health behaviors. A subsequent study by Smith and Waitzman (1994) found empirical support for this hypothesized synergy between low income and unmarried status among men. Using U.S. data from the National Health and Nutrition Examination Survey (NHANES) 1 Epidemiologic Follow-Up Study in 1982–1984, their results confirmed that unmarried men with low income exhibited all-cause mortality rates much greater than would be expected on the basis of income and unmarried status alone; yet, this pattern was not evident for women. Further, this pattern was more apparent for causes of death that have a behavioral component. Taken together, these two studies suggest that the combination of low income (or low education) and unmarried status may interact in a way that is deleterious to men's health in particular. Thus, relatively high mortality among unmarried men with low education may undergird gender differences in the overall educational gradient of mortality in the United States today.

The notion that a synergy between marital status and education on mortality risks exists for men but not for women is also indirectly supported by evidence of, and explanations for, gender differences in the health benefits of marriage. It is well documented that married adults experience lower mortality risks than unmarried adults. Further, several studies have reported that this disparity appears greater among men than women (Gardner & Oswald, 2004; Gove, 1973; Litwak & Messeri, 1989). One compelling explanation for the greater disparity among men points to health behaviors. Specifically, because men exhibit higher baseline levels of health-compromising behaviors than women, the social controls and spousal monitoring of health behaviors that often accompany marriage disproportionately benefit men (Umberson, 1987, 1992). Because unmarried men do not benefit from the social controls and shared lifestyle within marriage, their mortality risks are more highly dependent upon individual re-

sources such as educational attainment. Therefore, the educational gradient of mortality may be steeper for unmarried men than for married men but largely invariant to marital status among women. To be fair, selection processes may also contribute to the mortality advantage of married adults if healthier persons are more likely to marry and stay married. For example, adults who engage in risky health behaviors or exhibit physical characteristics associated with poor health have lower rates of marriage than healthier persons (Fu & Goldman, 1996; Murray, 2000). While these health characteristics appear to depress marriage rates similarly for men and women (Fu & Goldman, 1996; Gove, 1973), disadvantaged labor force characteristics depress rates more for men (Sweeney, 2002). Thus, unmarried men may be more negatively selected on employment characteristics than unmarried women, which could inflate the benefits of marriage among men.

As an alternative to the behavioral hypothesis described previously, Preston and Taubman (1994) hypothesized that men exhibit a steeper educational gradient of mortality because household well-being may be more closely linked to men's education than to women's. Indeed, women's mortality risks may have historically been more closely linked to their husbands' socioeconomic status than their own. For example, Arber (1987) showed that the gradient for chronic illness among married women was flatter when analyzing their own occupation, yet converged toward the steeper male gradient when they were analyzed using their husbands' occupation. The gendered nature of paid employment and family responsibilities does tend to constrain the lives of married women to be contingent on their husbands (Moen & Chermack, 2005), although this is less true today than it was just a generation ago.

Moreover, the increasingly central role that education has played in assortative mating in the United States (Schwartz & Mare, 2005) has implications for both hypotheses. First, assortative mating clouds the direction of causality for the behavioral hypothesis. As such, it is difficult to distinguish whether married men exhibit better health behaviors than unmarried men due to social controls and the shared lifestyle and resources within marriage or whether men with positive health behaviors are more likely to marry women with similarly positive health behaviors. Disentangling these two possibilities is not imperative for testing the hypothesis or for its explanatory utility, as long as this caveat is recognized. It does, however, provide additional motivation for evaluating gender differences in the gradient within the context of marriage. Assortative mating also suggests that these differences will be greatest outside of marriage. Second, the high degree of educational assortative mating exhibited today means it is increasingly unlikely that men's education is the dominant, upwardly driving force behind household well-being as proposed in the household hypothesis. In fact, recent research finds that the education of both spouses contributes to their mortality risks (Kravdal, 2008).

Table 1. Demographic Characteristics of Sample

Characteristic	Women					Men				
	All	Never Married	Married	Divorced or Separated	Widowed	All	Never Married	Married	Divorced or Separated	Widowed
Marital status (%)	100.0	4.0	57.0	7.9	31.1	100.0	3.8	83.2	5.9	7.1
Age at interview (years)	67.5	68.7	65.3	64.6	72.2	66.4	66.2	66.1	63.9	72.3
Education (%)										
Less than high school	30.8	25.5	26.0	26.9	41.1	31.3	37.4	30.0	33.2	42.5
High school	43.3	36.1	46.9	41.2	38.0	33.8	30.2	34.4	30.3	31.7
Some college	14.7	13.0	15.3	18.0	13.1	14.1	11.9	14.2	15.8	12.1
College	11.2	25.4	11.8	13.8	7.7	20.8	20.6	21.4	20.7	13.6
Number of deaths	31,325	1,448	13,750	2,211	13,916	31,733	1,394	25,077	1,930	3,332
Percent of deaths	31.3	36.8	24.1	28.5	44.5	39.6	45.4	37.5	41.5	59.0
N	100,004	3,939	57,037	7,766	31,262	80,204	3,070	66,833	4,652	5,649

Note: Demographics are weighted to reflect sampling design. Number of deaths and sample sizes are not weighted. The category "college" includes college degree or higher.

As noted earlier, we build on the work of Nathanson and Lopez (1987) and Smith and Waitzman (1994) by evaluating gender differences in the educational gradient of mortality within the context of marriage. We extend their work by using a newer data set that contains more recent cohorts and a much larger sample, although we diverge from these studies in two ways. First, we focus on adults 55 years of age and older because most deaths in the United States occur to this age range. Second, we select education as our measure of socioeconomic status because, compared with income and occupation, it is a more stable measure; it is available for men, women, and retired individuals; and it is more closely associated with health behaviors (Preston & Taubman, 1994; Winkleby, Jatulis, Frank, & Fortmann, 1992). Specifically, we use the combined 1986 through 1996 National Health Interview Survey Linked Mortality Files (NHIS-LMF) to examine the educational gradients of mortality between non-Hispanic White men and women 55 years of age and older. We first test whether the gender difference in gradient is statistically different when aggregating across all marital statuses and then test for differences within each marital status. Where we find significant gender differences, we examine leading causes of death to more systematically assess the behavioral hypothesis. We also evaluate whether men's steeper gradient is due to a stronger influence on household well-being by simulating whether the gradient among married women becomes steeper and converges toward the gradient for married men once we replace wives' education with their husbands' education.

METHODS

Data

Our data come from the public-use NHIS-LMF. The NHIS is a cross-sectional survey that has been conducted annually since 1957 and is the primary source of health information on the civilian, noninstitutionalized population of the United States (National Center for Health Statistics [NHCS], 2005). Through household interviews, the NHIS collects information on approximately 100,000 individuals each year. The National Death Index (NDI) is a computerized database of all certified deaths in the United States since 1979. The public-use NHIS-LMF links adult respondents in the NHIS to death records in the NDI through a probabilistic matching algorithm (Lochner, Hummer, Bartee, Wheatcroft, & Cox, 2008; NCHS, 2005). Our data contain NHIS years 1986 through 1996 with respondents linked to the NDI for mortality follow-up through December 31, 2002.

Sample

We selected non-Hispanic White adults between 55 and 84 years of age at the time of their NHIS survey, which corresponds to 55–100 years of age at death if the adult died during the follow-up period. We focused on non-Hispanic Whites because family structures (Spain & Bianchi, 1996) and educational gradients (Jemal et al., 2008) vary by race/ethnicity, which warrants separate analyses for each group. We restricted our sample to individuals aged 55–84 years at the time of survey for three reasons. First, 88% of deaths to non-Hispanic Whites occur after age 55 (Kung, Hoyert, Xu, & Murphy, 2005), ensuring that our findings are relevant for most deaths among this population. Second, because we were interested in the mortality of never married adults, establishing a lower limit of 55 years ensured a sufficiently long exposure to any lifestyle and risk factors associated with a never married status. Third, in preliminary analyses, we discovered that the matching of death certificates among women after ages 80–85 years at survey was less successful than it was for men, which could bias the results.

Among the 181,780 non-Hispanic White adults between 55 and 84 years of age interviewed between 1986 and 1996, we excluded 0.9% from further analysis because they were missing information on educational attainment or marital status. These selection criteria resulted in a final analytic sample of 180,208 adults, with 63,058 of them identified as subsequent deaths in the NDI. Table 1 contains key demographic information for the final analytic sample.

Methods

We estimated Cox proportional hazards models to test whether the educational gradients among men and women are statistically different. These models estimated the risks of death during the follow-up period for adults with less than a high school education, a high school education, or some college, using a college degree or higher as the reference group. The risk of death for individuals with a college degree or higher is, by definition, 1.0, and the risks of death for individuals with less than a college degree were anticipated to be greater than 1.0. The incremental increases in the risks between higher and lower education levels indicate the steepness of the educational gradient.

Marital status was assessed at the time of interview and was categorized as married, never married, divorced or separated, or widowed, as well as two aggregated categories that include the unmarried (never married, divorced, separated, or widowed) and the previously married (divorced, separated, or widowed). Model predictors include age, gender, education, and the interaction between gender and education. *Age* at the time of survey is a continuous measure ranging from 55 to 84 years. *Gender* is a dichotomous indicator using females as the reference group. We used a categorical specification of education defined earlier based on our model fit tests and the analysis of Backlund, Sorlie, and Johnson (1999), who found that a categorical specification exhibited a better model fit than a continuous specification for predicting mortality. A categorical specification also allows gender differences to exist along portions of the gradient. We compared model fits using the Bayesian information criterion (BIC). The BIC preferred the categorical specification when aggregating across all marital statuses, for married adults, and for unmarried adults as a group. Because others report similar (Zajacova & Hummer, under review) or better (Zajacova, 2006) model fit using a continuous specification, we also estimated a model using a continuous term for comparison.

For each marital status, we estimated three models. Model 1 includes the main effects of age at interview, gender, and education. Model 2 adds the education-by-gender interaction. A statistically significant interaction between education and gender indicates a gender difference in the educational gradients within a given marital status. Model 3 reestimates Model 2 using a continuous specification of education. All models were estimated with SUDAAN (2005) to account for sampling weights and survey design of the NHIS-LMF. Using the model diagnostics provided by SUDAAN, we manually calculated the BIC values to compare the fit of models using the categorical term with those using the continuous term (available on request). Lastly, we tested the proportionality of hazards assumption and found that although the hazards exhibited a marginal convergence with time, the effect was not statistically significant.

To evaluate the behavioral hypothesis, we estimated cause-specific proportional hazards models for the most common causes of death in the United States (Kung et al., 2005). These cause-specific death categories include the following underlying causes of death from the International Classification of Diseases (ICD)-10 113-group recodes: diseases of the heart (55–68), cancers excluding lung (20–43 except 27), lung cancer (27), cerebrovascular diseases (70), chronic lower respiratory diseases (83–86), accidental and violent deaths (114–129), diabetes mellitus (46), influenza and pneumonia (77–78), and chronic liver diseases and cirrhosis (94–95). We examined lung cancer separately because Nathanson and Lopez (1987) speculated that the relatively high prevalence of smoking among low-income men may contribute to the steeper gradient for men overall. The behavioral hypothesis is supported if the all-cause mortality models indicate that gender differences in the gradient are most pronounced among low-educated, unmarried adults and if the cause-specific models reveal a similar pattern primarily among causes with a strong behavioral component. An alternative, and perhaps more direct, approach to testing this hypothesis would control for cumulative exposure to certain health behaviors. However, the NHIS did not consistently collect in-depth health behavior data throughout the 1986–1996 period. Using causes of death as an indirect approach to test the behavioral hypothesis is not ideal, although it is one way to gain insights into cumulative exposures to certain health behaviors.

To evaluate the household hypothesis, we estimated three additional models that replaced married adults' own education levels with various measures of household education. The first model replaced women's education with their husbands' education, whereas the second replaced both spouses' education with their maximum education, and the third replaced both spouses' education with their median education. A similar approach has been used with occupational data (Krieger, Chen, & Selby, 1999). We estimated the median from our categorical specification of education (1 = less than high school; 2 = high school; 3 = some college; and 4 = college) such that a median of 1 indicates that both spouses had less than a high school education, a median of 1.5 indicates that one spouse had less than a high school education, whereas the other had a high school education, etc. We interpret our results as follows. If the educational gradient for married women is weaker than that for married men when modeling one's own education, yet becomes similar to men's when given their husband's education, then this supports the hypothesis that the health of married households is disproportionately influenced by men's education. Conversely, if those gradients diverge, or if the gradients for men and women become steeper and similar using the median education, then this contradicts the household hypothesis and instead supports the notion that the education of both spouses contributes to household well-being.

Table 2. Risks of Death During Follow-Up

	All Marital Statuses			Married			Unmarried[a]			Previously Married[a]			Never Married			Divorced/Separated			Widowed		
	1a	1b	1c	2a	2b	2c	3a	3b	3c	4a	4b	4c	5a	5b	5c	6a	6b	6c	7a	7b	7c
Age at interview	1.091**	1.091**	1.091**	1.094**	1.094**	1.094**	1.079**	1.079**	1.079**	1.079**	1.079**	1.078**	1.079**	1.079**	1.080**	1.082**	1.082**	1.081**	1.080**	1.080**	1.079**
Male	1.631**	1.510**	1.776**	1.721**	1.657**	1.731**	1.742**	1.616**	1.838**	1.736**	1.621**	1.790**	1.771**	1.620**	2.125**	1.824**	1.708**	1.966**	1.674**	1.601**	1.641**
Education (college)																					
Less than high school	1.565**	1.481**		1.605**	1.569**		1.444**	1.399**		1.420**	1.385**		1.540**	1.391**		1.589**	1.531**		1.373**	1.359**	
High school	1.297**	1.229**		1.328**	1.281**		1.231**	1.194**		1.212**	1.175		1.290**	1.298**		1.313**	1.254**		1.180**	1.159**	
Some college	1.210**	1.147**		1.222**	1.170**		1.170**	1.129**		1.149**	1.113**		1.259**	1.200†		1.219**	1.165†		1.122**	1.101*	
Education[b]			0.963**			0.956**			0.970**			0.969**			0.974**			0.959**			0.971**
Male × Education																					
Male × Less than high school		1.091**			1.028			1.088†			1.064			1.223†			1.066			1.029	
Male × High school		1.089**			1.050			1.082			1.098†			0.991			1.089			1.078	
Male × Some college		1.088**			1.062			1.104			1.095			1.117			1.085			1.075	
Male × Education[b]			0.992**			0.999			0.994			0.997			0.984†			0.993			1.001
–2LL[c]	2,360,705	2,360,694	2,360,825	1,427,301	1,427,299	1,427,382	850,899	850,896	850,943	745,093	745,090	745,125	88,735	88,729	88,745	132,295	132,294	132,295	593,507	593,505	593,535
Number of deaths	63,058	63,058	63,058	38,827	38,827	38,827	24,231	24,231	24,231	21,389	21,389	21,389	2,842	2,842	2,842	4,141	4,141	4,141	17,248	17,248	17,248
N	180,208	180,208	180,208	123,870	123,870	123,870	56,338	56,338	56,338	49,329	49,329	49,329	7,009	7,009	7,009	12,418	12,418	12,418	36,911	36,911	36,911

Notes: The category "college" includes college degree or higher. Hazard ratios estimated from Cox proportional hazards models.
[a] Unmarried includes never married, divorced/separated, and widowed. Previously married includes divorced/separated and widowed.
[b] Continuous measure of education used in Model c only. A risk less than 1.0 implies each year of education reduces the risk of death.
[c] –2 × log (likelihood)
†*p* ≤ .10; *p* ≤ .05; **p* ≤ .01.

RESULTS

Table 1 presents a summary of key demographic characteristics of our analytic sample. A few gender differences in the distributions of marital status and educational attainment are noteworthy. First, women were less likely to be married and more likely to be widowed. Just 57% of women were married compared with 83% of men, whereas 31% of women were widowed compared with just 7% of men. Although a similar proportion of men and women reported very low levels of education, almost twice as many men reported a college degree. Specifically, 31% of men and women reported having less than a high school diploma, whereas 21% of men and 11% of women reported having a college degree. The distribution of educational attainment within marital statuses reveals a sharp contrast between never married women and men in these cohorts. Never married women had the highest education levels compared with other women and with never married men, whereas never married men had relatively low education. In fact, never married women were more likely to have a college degree than any other combination of gender and marital status.

Table 2 reports the risks of death estimated from Cox proportional hazards models. Models 1a–c estimated the risks of death across all marital statuses. A value of 1.631 for men in Model 1a means that the risk of death for men was $100(1.631 - 1) = 63.1\%$ greater than the risk for women, the reference group. As expected, the risks were greater for men and for individuals without a college degree, and the risks increased with age at interview. Consistent with previous reports that the gradient is visually steeper for men than women, the education-by-gender interactions are statistically significant in Model 1b, and they are all greater than 1.0. For example, the risks of death for men with a high school diploma or less were roughly 9% greater than we would expect on the basis of gender and education alone.

Models 2b and 3b in Table 2 show the risks of death for married and unmarried adults, respectively. Among married adults, we found little evidence that the educational gradients of mortality were different between men and women. In contrast, we did find a marginal gender difference in the gradients among unmarried adults. Unmarried men with less than a high school education exhibited almost a 9% greater risk of death ($p = 0.07$) than would be expected on the basis of gender and education alone. Excluding never married adults from Model 3b, the model for previously married adults in 4b shows that these men with a high school education exhibited a slightly greater ($p = 0.10$) risk of death than expected. Gradient comparisons within specific unmarried statuses are shown in Models 5b, 6b, and 7b. The only specific unmarried status that retained an education-by-gender interaction is never married adults ($p = 0.08$), with men having less than a high school education experiencing a 22% greater risk of death than expected on the basis of gender and education alone. The relatively small sample sizes among the detailed
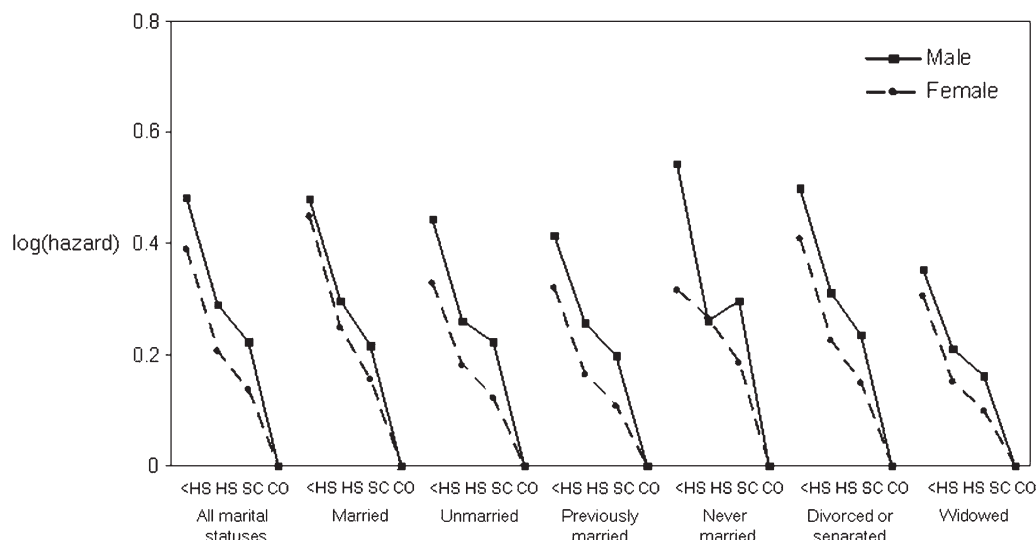
Figure 1. Educational gradients by gender and marital status. The log(hazard)s were estimated from Cox proportional hazards models stratified by marital status while controlling for age and education. Education categories are less than high school (<HS), high school (HS), some college (SC), and college degree or higher (CO) as the omitted reference.

statuses may preclude finding other differences, however. Figure 1 depicts the gradients.

In sum, we found that the gradient is steeper for men than for women when aggregating across marital statuses, although the difference is quite small. Table 2 and Figure 1 reveal that unmarried men were primarily responsible for the marginally steeper gradient. More specifically, never married men exhibited a steeper gradient along the primary and secondary education segment, whereas previously married men exhibited a slightly steeper gradient along the postsecondary segment. We should note that these conclusions were informed and validated by additional models (available on request) that used high school diploma as the reference group. The earlier results provide initial support for the behavioral hypothesis because gender differences in the gradients, albeit fairly small in size, are limited to unmarried adults. We now examine whether unmarried men were more likely to die from causes that are more behaviorally linked than their female peers.

Table 3 shows the risks of death estimated from nine cause-specific models for unmarried adults. The results for chronic lower respiratory diseases and lung cancer are the most consistent with those for all-cause mortality among unmarried adults in Table 2. Specifically, we found significant gender differences in the gradient for these two causes of death, with the most pronounced differences among adults with very low levels of education. In contrast to the magnitude of gender differences in the gradients for all-cause mortality, gender differences for smoking-related causes of death are impressive. In addition, unmarried men with less than a high school education experienced elevated risks of death from non-lung cancers and from influenza and pneumonia, although the small number of deaths from

influenza and pneumonia may hinder our ability to detect significant differences. Other causes of death with a strong behavioral link, including cirrhosis, diabetes, and accidental and violent deaths, did not exhibit a similarly stronger gradient for men. Taken together, these patterns support the behavioral hypothesis because the marginally steeper gradient among unmarried men, especially men with low education, was primarily responsible for the steeper gradient for men overall, with these men more likely to die from causes with a strong behavioral component.

Because we did not find significant gender differences in the gradient among married adults in Model 2b in Table 2, there is little evidence to support the household hypothesis. However, we continued with the proposed analysis to see whether, at least directionally, we might find marginal support for this hypothesis. Figure 2a–d displays the educational gradients for married men and women using their own education levels and three measures of household education estimated with the model form used in 2b from Table 2. Figure 2a shows the gradient using one's own education. Figure 2b displays the gradients when married women's education is replaced by their husbands' education. Because the women's gradient did not become steeper and converge toward the male gradient, we found little evidence that men's education was the primary determinant of household health. In fact, the strong departure from the original gradient in Figure 2a suggests that household health was influenced by both spouses' education. This interpretation is further supported by Figure 2c and d. Figure 2c displays marginally flatter gradients for both men and women when the maximum household education is modeled for each spouse, signaling that relevant information was lost in the model. Finally, the sharpest gradient appeared when the median household education was modeled

Table 3. Risks of Death Among Unmarried Adults From Common Underlying Causes

	Diseases of the Heart	Cancers (excluding lung)	Lung Cancer	Cerebrovascular Diseases	Chronic Lower Respiratory Diseases	Accidental and Violent Deaths	Diabetes Mellitus	Influenza and Pneumonia	Chronic Liver Disease and Cirrhosis
Age at interview	1.097**	1.048**	0.996	1.118**	1.047**	1.058**	1.041**	1.127**	1.012
Male	1.863**	1.302**	1.539*	1.105	1.000	3.334**	1.428	1.739*	7.269**
Education (college)									
Less than high school	1.577**	1.009	1.350*	1.138	1.691**	1.421	2.144**	0.984	3.038*
High school	1.136*	1.077	1.131	1.065	1.677**	1.197	1.778**	1.051	3.219*
Some college	1.135†	0.944	1.418**	1.084	1.328†	1.441	0.998	0.859	3.007*
Male × Education									
Male × Less than high school	0.958	1.268*	1.729**	1.143	2.200**	0.757	0.904	1.396	0.310†
Male × High school	1.089	1.130	1.452†	1.119	1.667†	1.049	0.871	0.994	0.229*
Male × Some college	1.026	1.172	1.226	0.828	1.723†	0.630	1.855	1.342	0.801
Number of deaths	8,781	3,661	1,674	1,926	1,531	476	648	841	177
N	56,338	56,338	56,338	56,338	56,338	56,338	56,338	56,338	56,338

Notes: The category "college" includes college degree or higher. Hazard ratios estimated from cause-specific Cox proportional hazards models. See Methods section for ICD codes.

†$p \leq .10$; *$p \leq .05$; **$p \leq .01$.

for each spouse, supporting the notion that education is a household resource to which both spouses contribute. The models behind Figure 2a–d are available on request.

DISCUSSION

Our findings for non-Hispanic White men and women aged 55 years and older in the 1986–2002 period confirm earlier reports that the educational gradient in mortality is visually steeper for men than for women. Furthermore, we found that this difference is statistically significant when aggregating across marital statuses, although the size of the difference is modest. Our analyses revealed that the marginally steeper gradient for men overall reflects a marginally steeper gradient among unmarried men compared with unmarried women. Never married men exhibited a steeper gradient along the primary and secondary education segment, whereas previously married men exhibited a slightly steeper gradient along the postsecondary segment. We did not find differences in the gradient between married men and women. Our results align with other research that found a synergistic effect between low income and unmarried status among working-age men (Smith & Waitzman, 1994) and with the hypothesis that the health behaviors and health
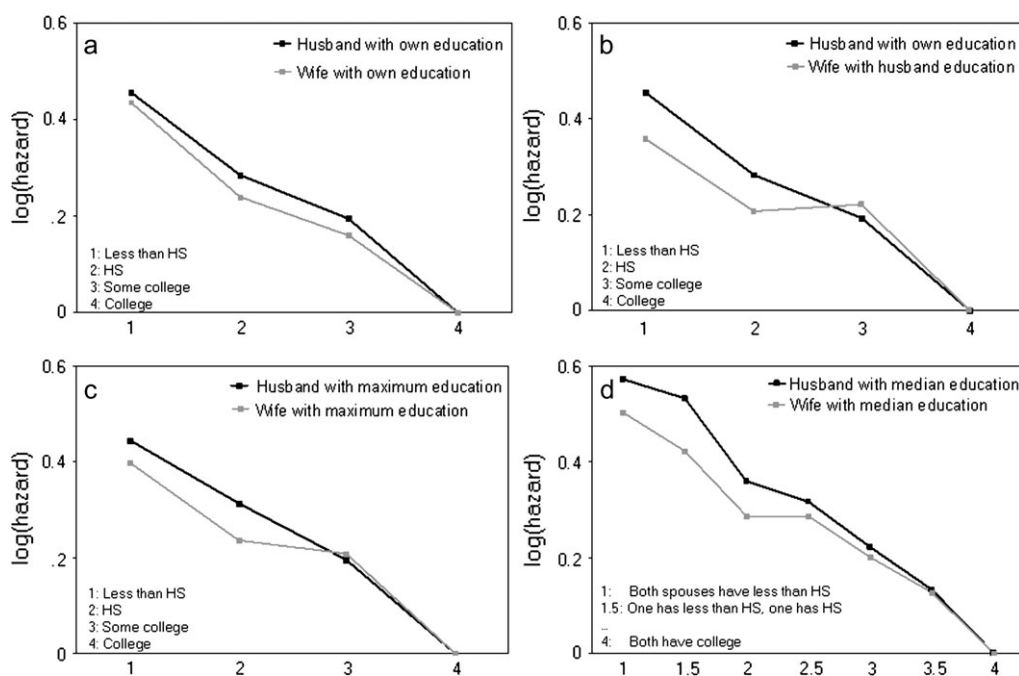

Figure 2. (a–d) Educational gradients among married adults using own education and three measures of household education.

status of men with low socioeconomic status are particularly vulnerable to a lack of health-enhancing social ties such as marriage (Nathanson & Lopez, 1987).

Our tests of two hypotheses to explain the marginally steeper gradient among men provided strongest support for the behavioral hypothesis. Gender differences in smoking patterns seem to be the most relevant contributor because deaths due to lung cancer and chronic lower respiratory diseases exhibited a strong education-by-gender interaction. Other causes of death with a strong behavioral etiology, including cirrhosis, accidental and violent deaths, and diabetes, did not exhibit a similar pattern. In sum, our finding that differences in the gradient were limited to unmarried adults, combined with the finding that unmarried men with less than a high school education were more likely than their female peers to die from causes for which smoking is a major risk factor, supports the behavioral explanation proposed by Nathanson and Lopez (1987). We should note that additional analyses (available on request) indicated that low-educated, married men were also more likely to die from lung cancer than their female peers, although gender differences in smoking-related deaths were more pronounced among unmarried adults.

It is well documented that both informal (e.g., marriage) and formal (e.g., workplace, medical care) social supports reduce mortality risks (Litwak & Messeri, 1989). Unmarried, low-educated men do not benefit from marriage as a key informal support, and they likely have less access to high-quality formal supports due to their low education and compromised occupational opportunities. Although their female counterparts may also have less access to formal supports, women tend to have larger and closer personal networks than men (McPherson, Smith-Lovin, & Brashears, 2006), which may partly compensate for a lack of spousal support. In fact, unmarried women report more attempts by others to regulate their health behaviors than do unmarried men, and this gender gap is most pronounced among the never married (Umberson, 1992). Similarly, we found the greatest gender disparity in the gradient among never married adults.

Although our results support the behavioral hypothesis for non-Hispanic White adults 55 years of age and older, they may not apply to other cohorts or race/ethnic groups that tend to exhibit different health behaviors and/or social supports. For example, across lower levels of education, the gradient in life expectancy at midlife appears weaker for Hispanic men compared with non-Hispanic White men; yet, this discrepancy is not found among their female counterparts (Lin et al., 2003). This might explain why a recent study that combined these two ethnic groups did not find gender differences in the educational gradient of mortality (Turra & Goldman, 2007) because the steeper gradient among low-educated, non-Hispanic White men that we found may have been counterbalanced by a weaker gradient among Hispanic men, thereby generating an average male gradient that was similar to the average female gradient.

We found little evidence to support the household hypothesis that posits that educational gradients in mortality are steeper for men because household health may be disproportionately influenced by men's education levels. In fact, the gradients between married men and women in our sample were remarkably similar, corroborating a previous U.S. study (Zajacova, 2006). Our simulation of the gradients using various measures of household education in lieu of individual education suggests that mortality risks are likely shaped by the combined education of the household in addition to one's own education. This accords with research using occupational status (Krieger et al., 1999), with research that demonstrates that the education of both spouses influences their mortality risks regardless of gender (Kravdal, 2008), and with the notion that resource pooling, broadly defined, among spouses inextricably links their health experience (Lindau, Laumann, Levinson, & Waite, 2003). Although we found little support for this hypothesis in our cohorts who exhibited a high degree of assortative mating, it may hold for others with large, systematic disparities in education or occupational statuses between spouses (Arber, 1987). Further, these results may not apply to younger cohorts where educational gradients of mortality are generally steeper, and marriage benefits for health outcomes are generally larger.

In ancillary analyses (available on request), we examined whether greater financial returns to education for unmarried men compared with unmarried women might also contribute to their slightly steeper gradient. We found little support for this alternative explanation. Controlling for income did not materially change the magnitude or significance of the education-by-gender interactions for unmarried adults, which concurs with other studies of gender differences in the educational gradient of mortality (Zajacova, 2006) and depression (Ross & Mirowsky, 2006), but not all (Elo & Preston, 1996). However, an economic explanation should be fully vetted using a data set that contains appropriate economic measures for this age group, such as wealth, before it is dismissed. The NHIS-LMF does not collect information on wealth, and the detailed income measure it does provide is top coded at $50,000 with roughly one quarter of our sample not providing these data. Thus, our ancillary analyses relied on a summary measure that was available for most of our sample that simply indicated whether or not family income was above $20,000. We also tested whether controlling for self-rated health at the time of interview would diminish the main (Feinglass et al., 2007) and interactive effects of education. This control attenuated the main effects somewhat; yet, it did not influence gender differences in the gradient among unmarried adults.

Our study has several strengths, including its recent and large nationally representative sample, as well as information on causes of death as indicators of risk factors. A few limitations must be noted, however. First, marital status is assessed once at the time of survey. We set the lower age

limit at 55 years at the time of survey to minimize the potential for subsequent transitions into (re)marriage and to allow us to detect the effect that at least 55 years of being never married has on mortality. A second limitation is that we cannot rule out selection effects. The low-educated, never married men in our sample may have experienced a lifetime of poor health, making them undesirable spouses and accelerating their mortality risks regardless of compromising health behaviors. Third, we measure only one type of informal social support. Thus, we cannot account for other informal sources such as parents, children, extended family, friends, or colleagues. Another potential limitation is our use of cause of death information as an indirect test of the behavioral hypothesis. We hope that our findings launch future work using detailed health behavior data. Finally, although the algorithm used to link NHIS survey records with NDI death records is reported to correctly classify the vital status of 98.5% of survey records (NCHS, 2005), the accuracy of the underlying cause of death information on death certificates is a widely recognized concern (Sehdev & Hutchins, 2001).

Taken together, our results present new and important findings about the educational gradients in mortality among men and women. As an increasing proportion of men approach their retirement years with a checkered marital history, the role of education in reducing their mortality risks is perhaps more important than ever. Future research should monitor these gradients to see whether and how the confluence of gender-specific trends in smoking behavior, the narrowing gender gap in mortality, and the growing education gap in mortality, which has recently grown more for women than for men (Meara, Richards, & Cutler, 2008), may redefine gender differences in the gradient. Future research should also evaluate these gradients for other race and ethnic groups, as well as working-age adults, because the patterns we found may not apply to these populations.

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