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Gender and the Socioeconomic Gradient in Mortality*

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Despite considerable evidence documenting a strong and persistent relationship between socioeconomic position and mortality, recent research suggests that this association may be weaker among women. In our examination of gender differences in the socioeconomic gradient in mortality, we argue that this inconsistency arises from the failure to consider the ways in which gender is a fundamental constituent of socioeconomic position. The data used are from the Panel Study of Income Dynamics. Respondents, including all household heads and their partners, aged 29 years and older in 1972 (N = 5,665; 56% female), were followed until 1991, death, or attrition. Discrete time event history analysis was used to examine the predictors of death between 1972 and 1991. Of the key socioeconomic predictors, years of education was measured at baseline, while earned income was a time-varying covariate. We find no gender differences in the effect of respondents' own socioeconomic positions on their mortality risk. However, increasing spousal income raises men's odds of dying, while the opposite is true for women. Our results raise questions about the prevailing view that the socioeconomic gradient in mortality is weaker among women. Moreover, gender differences in the effects of spousal earnings on mortality risk suggest that their labor market rewards have fundamentally different meanings for women and men.

In one of the earliest sociological treatises on the association between social class and health, Frederick Engels ([1844] 1984) exposed the ravages of industrialization in the abysmal living and working conditions that shortened the lives of “labourers, mechanics and servants,” compared with the “gentry and professional persons.” Documentation of the health consequences of socioeconomic inequality has continued since his observations were recorded over a century ago. This association has become a key tenet of medical sociology and social epidemiology: that differential exposure to socioeconomic resources defines individuals’ general susceptibility to disease and early death (Gerhardt 1989). Recent observations that socioeconomic
inequalities in women’s health are “unsystematic, inconsistent or insignificant” (Dahl 1991: 499) raise questions about the generalizability of this proposition. Does the unequal distribution of material and social rewards stemming from systematic structures of inequality really not affect the health of women? Or, does the ambiguity arise from failure to consider the ways in which gender is a key constituent of socioeconomic position?

Women are largely invisible in classical theories of social stratification. According to such perspectives, the work position of male heads determines the social class of all household members, while women’s activities and work lives are deemed to be irrelevant to processes of class formation and class action (Sorensen 1994). Feminist theory and research has long rejected this view, asserting that the socioeconomic positions of women and men emerge from social relations of production (and consumption) that are inextricably linked to those operating with the family.

This paper explores relationships among gender, socioeconomic position, and mortality from the premise that individuals’ structural positions and their associations with mortality reflect specific conditions and processes in the labor market and in the family that are different for women and men. We begin with a discussion of the problems inherent in conceptualizing and measuring women’s socioeconomic position, and we turn to research investigating the socioeconomic patterning of mortality by gender. Our own examination of this relationship suggests that restricting the conceptualization and measurement of socioeconomic position to that of male household heads will fail to detect mortality effects among women that emerge from their own positions in the labor market, as well as those among men that stem from the socioeconomic positions of their spouses.

SOCIOECONOMIC POSITION AND WOMEN

Stratification theorists operate under the broad assumption that a complex of social institutions generates economic and social inequalities in access to the assets, resources, and valued goods that underlie stratification systems (Gruyter 1994). Those of a classical orientation contend further that stratification structures and processes are universally represented by the experiences of the male household head. In a heated debate in the mid-1980s, Goldthorpe (1983:468) argued that a woman’s socioeconomic position is most adequately reflected by the occupational class of her husband because men have a “directly determined position within the class structure” stemming from their more enduring commitment to the labor market. Conceptually, this “conventional” view implies the existence of a family-based class1 position. Empirically, it means assigning women to the occupational class of male relatives.

Stanworth (1984) and others (Abbott 1987; Leifurfrud and Woodward 1987; Crompton and Mann 1986) countered that the conventional approach was not credible because women’s increasing participation in the labor force exposed them to the same forces of social cleavage as men. Many also argued that such processes and their consequences could only be understood by considering gender as integral in shaping asymmetrical relations of power. The research implications of this alternative, are that women’s socioeconomic positions are not necessarily reflected by those of the men they live with and, importantly, that women’s labor market experiences affect not only their own lifestyle, standard of living, and life chances but also those of all household members. Operationalizing this theoretical perspective entails using individual-based socioeconomic characteristics (Health and Britten 1984; Abbot 1987; Stanworth 1984) or reinterpreting “family” socioeconomic position to encompass the characteristics of both husbands and wives (Bonney 1988; Graetz 1991; Hayes and Jones 1992). This debate made women more visible in theory and research on social stratification and social class, but, as the following section argues, the gendered nature of differential access to productive resources and other forms of power that constitute socioeconomic position has not been taken up in a systematic way in studies of its mortality consequences.

SOCIOECONOMIC POSITION, GENDER, AND MORTALITY

A substantial and growing body of research demonstrates a persistent, inverse relationship between socioeconomic position and health
(for reviews, see Feinstein 1993; Williams and Collins 1995; Bunker, Gomby, and Kehrer 1989). Some of the clearest evidence in the United States comes from studies of mortality. An early ground-breaking investigation that linked 1960 death certificates to socioeconomic data from the census of the same year showed mortality increasing as education and family income declined (Kitagawa and Hauser 1973). Those with very little formal schooling (0–4 years) were particularly vulnerable. More recent work using other nationally-based data sets suggests that these mortality differentials have endured over time (Feldman et al. 1989; Rogot, Sorlie, and Johnson 1992; Elo and Preston 1996; Pappas et al. 1993) and across a range of indicators of socioeconomic position. For example, with data from the National Longitudinal Mortality Study, Sorlie, Backlund, and Keller (1995) demonstrated higher odds of dying among laborers and technical, production, and service workers, compared with their professional and managerial counterparts. The risk of mortality is even greater the longer one stays in a low status job (Rogers and Carrigan 1996). Living in households with inadequate incomes (Zick and Smith 1991; Kaplan et al. 1987) or persistent low income (McDonough et al. 1997), or in neighborhoods where a large proportion of inhabitants has low income and little education (LeClerc, Rogers, and Peters 1997; Kaplan et al. 1996; Haan, Kaplan, and Cartaacho 1987) also increases the risk of dying.

Despite this apparent overwhelming evidence, socioeconomic inequalities in women's mortality are less consistent (Macintyre and Hunt 1997). With some exceptions (e.g., Martikainen 1995), most studies report stronger socioeconomic effects for men's mortality (Feldman et al. 1989; Lahelma and Valkonen 1990; Elo and Preston 1996; Pappas et al. 1993; Koskinen and Martelin 1994), even though formal tests of gender differences in this relationship are rarely done. These patterns apply to both individual and household measures of socioeconomic position, with the latter most typically represented by husbands' characteristics (Krieger, Williams, and Moss 1997).

One reason for the weaker effects of socioeconomic position on women's mortality risk is that studies fail to consider the ways in which socioeconomic position is gendered. For example, occupation is constructed from classification schemes that usually employ prestige, education, or income characteristics of occupations of males in the labor force (Haug 1977). Gender-based differences in income and education within occupations as well as different patterns of occupational concentration and socioeconomic desirability raise questions about the extent to which such classification systems are representative of an occupational hierarchy of women's jobs (Arber 1997). Moreover, an occupation-based measure of social class leaves those unwilling to assign women to the occupational group of their husbands without a system for categorizing women who are not in the paid labor force.

Some advocate the use of education as an alternative indicator since it can be assessed for everyone. The "human capital" dimensions of education that link it to health are apparent in its close connections to work and economic circumstances, psychosocial resources, and health behaviors (Ross and Wu 1995). However, gender differences across these diverse experiences suggest that education is not gender neutral. For example, in a study using Michigan death certificates from 1989 through 1991, Christenson and Johnson (1995) reported that secondary education (compared with only primary schooling) lowered women's mortality rates more than it did for men, but men benefitted more from post-secondary schooling (compared with only secondary education). They attributed the latter difference to the higher earnings of well-educated men that allow them to invest in healthy lifestyles. Other data support the notion that education is gendered. At least two studies report the narrowing of gender differences in the effect of education on mortality risk once adjustments were made for household income, occupation, employment status, and a number of socio-demographic characteristics (Sorlie et al. 1995; Martelin 1994). The change occurred mainly through reductions in men's relative risk of death, suggesting that the significance or meaning of education as it pertains to mortality risk may be different for men and women.

The same may be true for asset-based measures such as household income. Implicit in research examining this resource is the assumption that it is equally shared by all household members (Backlund, Sorlie, and Johnson 1996; Elo and Preston 1996; Martelin 1994; Pappas et al. 1993). However, research
on distributive relations within households suggests that power over financial decision-making and access to resources is gendered in ways that raise men’s standards of living relative to those of other unit members (Volger and Pahl 1993, 1994). For example, while men generally contribute more money to the household than women (in absolute terms), they hold back more for personal use (Pahl 1990). In a study of child benefits in the United Kingdom, researchers found that women who received cash payments directly spent a larger proportion of transfer allowances on clothing for their children and themselves than when payment was in the form of tax credits on the earned incomes of their husbands (Lundberg, Pollack, and Wales 1997).

A final issue that may account for research inconsistencies in gender inequalities in mortality is the temporal measurement of socioeconomic position. Most studies use cross-sectional measures, typically collecting them either concurrently with health status or at a more distal point in time relative to vital status assessment (i.e., baseline measures) (Bucher and Ragland 1995; Duleep 1986; Sorensen et al. 1995). While this may be appropriate for education which tends to be well-established by adulthood, occupation may change, and income may be quite volatile throughout the life cycle. Prime-aged women (i.e., 25–55 years), for instance, are more likely to experience income loss over time than their male counterparts, and this is frequently associated with marital separation or divorce (Duncan 1988). These observations suggest that, more so than is the case for men, baseline income may not accurately reflect women’s socioeconomic circumstances at the time of death.

RESEARCH QUESTION AND HYPOTHESES

Using data from the Panel Study of Income Dynamics (PSID), we examine gender differences in the relationship between socioeconomic position and mortality by incorporating prior theory and research on stratification and gender in ways that have been done in few studies to date. We argue that socioeconomic position (and its effect on mortality) is constituted by gender-based social relations in the “productive” sphere and in the family. In this regard, we consider measures of socioeconomic position that represent individuals’ labor market resources and rewards, namely education and earned income. We also examine representations of spousal education and earned income that shape the economic and social environment of households and their members. Longitudinal measures allow us to more accurately examine individuals’ socioeconomic positions in the context of social relations that may change over time.

We hypothesize that respondents’ education will have a stronger impact on men’s risks of dying than it will on women’s. However, because of men’s greater economic return on education, we expect that the addition of earned income to analytic models will attenuate or even eliminate any gender differential in education’s effect on mortality risk. The impact of spousal education may be more complex. For example, one of the ways in which men’s health is enhanced by marriage is from the greater tendency of women to influence the health behaviors of their partners (Umberson 1992). Because of the strong association between education and health behaviors (Winkleby et al. 1992; Shea et al. 1991), we hypothesize that wives’ education (as a proxy for preventive health practices) will be more important for the mortality risks of their husbands than husbands’ education will be for that of their wives.

Finally, to the extent that it can be assumed that individuals exert control over their own earnings, we anticipate finding no gender differences in the impact of respondents’ earned income on the risk of mortality. Expectations for the effect of spousal earned income on mortality risk are less clear. The null hypothesis of no gender difference in this relationship is possible if the health returns of spousal earnings are equally distributed within households. However, at least two conditional effects of spousal earnings also present themselves. First, because men’s average incomes are higher, they may be a better reflection of the lifestyle and life chances of household members. If this is true, spousal earnings will be inversely related to women’s mortality risks but will have a minimal impact for men. In contrast to the latter contention, research documenting an inverse relationship between men’s psychological health and the income of their spouses (Rosenfield 1992) suggests that men’s risk of mortality may increase along with the earnings of their wives because
women's earnings constitute a threat to men's perceptions of their role as chief breadwinner.

METHODS

Data

The PSID is an on-going longitudinal study of a representative sample of men, women, and children living in the United States and of the family units in which they reside (Hill 1992). Starting with a national sample of nearly 5,000 households in 1968, 49 percent of the individuals originally interviewed in that year were still in the study by 1989 (Fitzgerald, Gottschalk, and Moffitt 1998).\(^2\) The sharpest annual sample loss occurred between 1968 and 1969 when 12 percent of the sample left the study, but annual attrition rates since then are between 2.5 percent and 3 percent.\(^3\)

All household heads (defined as husbands in married-couple households or either a male or female in single-head households) and wives still in the PSID in 1972 and 29 years of age or older constitute the study sample (N = 5,665). Death or attrition withholding, individuals were interviewed annually between 1972 and 1991. Hence, the analysis examines survival patterns over a twenty year period from 1972 to 1991. There were 1,461 deaths (unweighted), 48 percent of which were among women.

Measures

Mortality. The event under investigation is all-cause mortality,\(^4\) a dummy variable coded 1 if a respondent died between 1972 and 1991 and 0 otherwise. Death is recorded as a reason for attrition from the sample. In the majority of instances, deaths are reported at the next annual interview by a surviving household member. For persons who were living alone when last interviewed, information about death may come from a surviving contact person, the administrator of the deceased person's estate, or the post office via returned mail.\(^5\) The most difficult, but rare, cases are those in which the individual left the study and neither other household members nor contact persons are available. In some cases, death information is obtained from family members residing in other PSID study households. In others, the fact of death cannot be determined without additional investigation.\(^6\)

Socioeconomic position. Education and earned income were used as indicators of socioeconomic position. Education is measured as years of formal schooling completed by 1972 (range: 1–24) and centered on the sample mean. Earned income is a five year average of individual annual earnings from wages, bonuses, overtime, commissions, and a professional practice or trade, and the labor portion of farm, market gardening, room and board, and business income. Reported for household heads and their spouses, it is measured in thousands of dollars, inflated to 1993 dollars and centered on the sample mean.\(^7\)

Sex, marital status, and controls. Sex is coded 1 for women and 0 for men. Marital status is represented by a dummy variable that codes those currently married or cohabiting as 1 and all others as 0. All analyses include controls for race, age, and work disability. Race was assessed by interviewer observation and is a dummy variable contrasting blacks (coded 1) with nonblacks (coded 0). Work disability, a lagged variable reported four years prior to the ascertainment of vital status, allows us to control, to some extent, for health selection into various income groups. Such selection could result in spurious associations between income and mortality because illness may result in both low income and death. Heads of households are asked annually: "Do you have any physical or nervous condition that limits the type of work or the amount of work you can do?" Identical information about wives' work disability was obtained only sporadically until 1981 when it became a stable feature of the interview. Work disability is represented by a dummy variable coded 1 for the presence of disability and 0 otherwise. Respondents who failed to answer the disability item or who were not heads of households during the years when disability information was gathered solely for heads were combined and a dummy variable representing "missing" on work disability was created. Respondents without disability information for a given year are coded 1 and all others 0.

Analytic Model

Discrete time event history analysis was used to examine whether and when death
occurs across the interview years. The hazard of dying, \( h(t) \), is the probability that death will occur in year \( t \), given survival until the beginning of year \( t \). This conditional probability is a function of time-varying and fixed covariates and can be represented as a logit function:

\[
\text{logit} \ h(t) = XA + Z(t)B,
\]

where \( X \) is a vector of fixed covariates, \( Z(t) \) is a vector of time-varying covariates and \( A \) and \( B \) represent vectors of parameters to be estimated.

Estimation of the mortality hazard was done using a file that pools together person-years of observation. This file contains 42,921 observations for women and 31,189 observations for men. Although using person-years rather than individuals as the unit of analysis inflates the number of observations in the file, estimates of standard errors and statistical significance tests are provided in an appropriate manner (Petersen 1986, 1991).

Each person-year record includes information on whether or not the person died in that year \( t \) and on time-varying and time-independent covariates. With the exception of race and education, all predictor variables are time-varying. Income is a lagged average of annual income ascertained over the five years prior to the year in which vital status was determined \( (t_{5-5}) \). As noted, disability status was measured four years prior to the event measurement \( (t_{4-1}) \), while age and marital status were determined at \( t \).

In examining gender differences in the effects of education and income on the risk of mortality, individual and spousal indicators of these constructs were used. The effects of spousal measures are, of course, relevant only to those who are married. Ross and Mirowsky (1992) elaborate a method initially proposed by Cohen (1968) that estimates the effects of conditionally relevant variables while simultaneously including those who are married and nonmarried in the estimates (see also Avison 1995). Essentially, the model compares those who are married with those who are not, while estimating the effects of variables that apply only to the married. For example, the effects of the respondent’s education \( (E_R) \), marital status \( (MS) \), and spouse’s education \( (E_S) \) on the risk of mortality are estimated by the following:

\[
\text{logit} \ h(t) = b_1(E_R - \bar{E}_R) + b_2MS + b_3(E_S - \bar{E}_S)MS
\]

The hazard of dying is regressed on respondent’s education (measured as deviations from the mean), marital status (a dummy variable where 1 equals married or cohabiting and 0 equals nonmarried), and the product of marital status and the mean deviate of spouse’s education. For those who are not married \( (MS = 0) \), the equation reduces to:

\[
\text{logit} \ h(t) = [b_1(E_R - \bar{E}_R)]
\]

Among those who are married, the expression becomes:

\[
\text{logit} \ h(t) = [b_1(E_R - \bar{E}_R)] + [b_2 + b_3(E_S - \bar{E}_S)]
\]

If Equation (3) is subtracted from Equation (4), the difference in the likelihood of dying between the married and the nonmarried depends on the educational qualifications of the spouse:

\[
\text{logit} \ h(t) = [b_2 + b_3(E_S - \bar{E}_S)]
\]

According to this formulation, those who are married differ from the nonmarried by an average amount (represented by \( b_2 \)) plus a deviation that depends on spousal education levels (represented by \( b_3 \)) (Ross and Mirowsky 1992:224). In the analyses that follow, spouse’s education and income are computed as conditionally relevant variables.

Sampling weights were used in all estimations to adjust for differential initial selection probabilities, mortality, and nonresponse (Institute for Social Research 1992). Because the weights differ across the time period considered here, time-varying weights representing the weights for each interview year are used in the analyses. Additionally, the effect of a complex sampling design on variance estimates was taken into consideration by using the jack-knife repeated replication method to calculate standard errors (Wolter 1983).
RESULTS

Sample Description

Descriptive statistics for the sample as it appeared in 1972 (baseline sample) and after it was converted into person-years of observation are presented in Table 1 (education and income are uncentered for descriptive purposes). The mean age of the baseline sample and average education of both respondents and their spouses was similar across sex, but men's average five-year earned income was almost five times greater. The average baseline earnings of all women are about $1,200 less than those for married women (see the earned incomes of men's spouses). This likely reflects the older average age of the nonmarried group (57 years compared with 47 years for married women) who constitute one-third of the total sample of women and are more likely to be retired. Men were more likely to be married, while black men comprised a slightly smaller proportion of the male baseline sample than did black women of the 1972 female sample.

The aging process is evident in several characteristics of the person-years sample. Average age, allowed to vary over time, increased by just over seven years for women and six years for men (women's slightly older average age may indicate their mortality advantage), and the slight decrease in men's average earned incomes may reflect the growing proportional contribution of retirees. A smaller percentage of the sample was married, while a larger proportion reported work disability.

Interestingly, in contrast to the pattern for men, average income was higher for the female person-years file than it was in the baseline sample. Average income for all

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*There were 1,875 married women and 1,972 married men in the sample derived in 1972.

*The baseline N (5,098) is smaller than the figure given in the text because of list-wise deletion of observations with missing data across the study variables.

*There were 25,270 person-years of observations among married women and 26,858 among married men in the analytic sample.

*N derived after list-wise deletion of observations with missing data across the study variables.
women rose by 11 percent, while that for married women increased by 24 percent. This could reflect a number of social processes, such as the narrowing income gap between men and women and selective attrition and mortality patterns that result in socioeconomically advantaged women making a larger proportional contribution to the total person-years of observation. The latter may also account for the slight increase in average years of education in the person-years sample.

Respondent's Education

The models presented in Table 2 examine the central hypotheses of the study. Model 1 tests whether there is an education gradient in the odds of dying among respondents over the study period, controlling for gender, age, race, and work disability. The observed negative coefficient for individuals' education is expressed as an odds ratio less than 1 and indicates that a year's increase in education (above the sample mean of 0) significantly lowers the mortality hazard by 3 percent. The odds of dying among women were half those of men (OR = .51), while individuals reporting a work disability or who had missing information on this item experienced increased odds of dying, compared with their counterparts reporting no

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N = 75,110

* p < .01; ** p < .001

*Measured as deviations from the mean for the whole sample.

*Measured as deviations from the mean for married persons.
disability. Not surprisingly, age and the hazard of mortality were positively associated, but no significant race differences in this outcome were detected.

Model 2 examines whether there are gender differences in the effect of the respondent's education on mortality by adding a multiplicative interaction term involving sex and respondent's education. The nonsignificance of the interaction indicates that respondent's education had the same effect on the likelihood of dying for women and men. This is contrary to our hypothesis that education would exert a stronger effect on men's risk of dying until controls for earned income were introduced.

Respondent and Spousal Measures of Income and Education

Model 3 adds average five-year annual income to the regression equation to examine the effects of respondents' socioeconomic characteristics on the odds of dying, net of one another. For those who are married, it also tests whether spousal earnings and education contribute to these predicted odds. The effect of respondent's education on mortality risk was reduced to nonsignificance with these additions. Further analyses revealed that earned income was more important in attenuating the effect of respondent's education on the hazard of mortality for men than it was for women, but spousal education was more consequential than earned income for both women and men (results available on request). Although the spousal education effect was of borderline significance in the pooled sample (OR = .97, t = -1.75), these findings suggest that respondent's education gets converted into "health capital" through returns on earnings and assortative mating processes that permit access to spousal economic and social resources. At least one study supports this contention in the Finnish context (Martikainen 1995), but another analysis of PSID data found no evidence of a spousal education effect on mortality risk among married persons (Smith and Zick 1994).

Unlike education, personal earnings exerted a significant effect on the mortality hazard that was independent of other variables in the model. A unit increase in respondent's earnings (i.e., $1,000 above the sample mean of 0), lowered the likelihood of dying by 1 percent over followup. The odds of dying for married individuals were lower than those of the nonmarried (OR = .71), but model 3 suggests no additional effect of spousal earnings. The latter observation changed, however, when gender differences in spousal socioeconomic characteristics were considered.

Model 4 tests whether the effect of personal earnings, spousal earnings, and spousal education on the mortality hazard are different for women and men. A statistically nonsignificant interaction between sex and personal earnings (OR = .99, t = -1.01) indicates that this was not the case for the latter variable. A nonsignificant estimate for the married coefficient suggests that the likelihood of dying among married individuals did not differ from that of the nonmarried at sample mean levels of spousal socioeconomic characteristics. However, a significant interaction between gender and spouse's earned income indicates that the odds of death for the married deviated from nonmarried levels as spousal earnings fell above or below the sample mean and, further, that the nature of this deviation depended on gender. Figure 1 illustrates this relationship. The fitted odds of dying are displayed separately for married men and women across spousal earnings that are allowed to range from the sample mean of 0 to three standard deviations above and below the mean. All other variables in the model are fixed at their observed sample mean or proportion values (Fox 1987). Women's risk of dying decreased as spousal income increased, but the opposite was true for men (both slopes are statistically significant). This finding accounts for the nonsignificant effect of spousal income observed in model 3 where its averaging across genders canceled out the separate effects of opposite sign for women and men. In contrast to spouse's earnings, there were no significant sex differences in the impact of spousal education on mortality, although a marginal inverse effect of this predictor on mortality is evident (OR = .96, t = -1.85).

DISCUSSION

The study was centrally concerned with relationships among gender, socioeconomic position, and mortality. We found the effects of respondents' own education and personal earnings on their risk of dying to be the same for women and men. Both socioeconomic characteristics were inversely related to mortality.

Note: Fitted odds of dying are calculated at spousal income values up to three standard deviations above and below the sample mean (0) and the sample means for all other independent variables (see Table 2, model 4).

These findings suggest that, at least at one extreme endpoint of the health continuum (i.e., death), experiences of structural inequality may not be embodied in different ways for men and women. Although the results cannot specify the nature of the links between socioeconomic position and mortality, they do suggest that the structures and processes that these positions imply have similar mortality consequences for men and women.

Having said this, our second key finding of gender differences in mortality risk among the married when spousal earnings are taken into consideration suggests that gender as an analytic category signifying different experiences should not be abandoned. Notably, women's earned incomes do matter to their husbands' risks of dying, but in a way that is fundamentally different from the effect of men's incomes on the mortality risks of their wives. Increasing spousal income is an asset for women but a liability for men. To our knowledge, this finding has not been reported elsewhere for mortality, although Rosenfield (1992) documented an inverse association between men's psychological health and their wives' relative incomes that was especially pronounced at the highest levels of family income. She explained her findings with reference to husbands' loss of status and power when their wives contribute significantly to family income. Higher levels of market work (reflected in higher earned incomes) may also leave women with less time for care-taking activities that enhance their husbands' health. In fact, Kessler and McRae (1982) found that women's employment was detrimental to the psychological well-being of their male partners, but, unlike the findings of the current study, men's reports of depression and low self-esteem were inversely related to their wives' incomes.

Of course, spurious and artifactual results cannot be ruled out in the current study. Health
selection may account for the positive association between men's risk of dying and the incomes of their wives. Important, it may reflect the necessity of wives' greater economic contributions because of their husbands' poor health, but the inclusion of baseline disability in the models controlled for this possibility to some extent.

Our finding of no gender differences in respondents' socioeconomic economic positions on mortality risk seems to be at odds with several other studies that report stronger effects among men. Almost without exception (Laelina and Valkonen 1990), categorical measures of education were used in other work, in contrast to the continuous form employed in our analysis. It may be that gender differences in education effects on mortality are more sensitive to qualitative differences in education expressed as completion of primary or secondary school or the possession of a college degree, but conversion of our measure into a categorical format failed to detect such differences (analyses not shown here). Nevertheless, it is difficult to compare work that varies considerably in relation to historical period and national jurisdiction.

Studies of income and mortality rarely explicitly specify the effects of spousal earned income on the mortality risks of both women and men. The ways in which this and other resources get distributed within households would be helpful in future research, especially in light of the contention that households are sites of conflicting individual interests expressed through differential power over decision-making and access to resources (Arber 1993). Similarly, additional work could focus on developing other measures of socioeconomic position that explicitly incorporate gender divisions in their formulation. For example, in work on the gender gap in pay, Paula England and colleagues (England 1992; England et al. 1994) differentiate occupations on the basis of cognitive skill and training demands, social skill demands, physical skill demands, amenities, disamenities, and industrial and organizational characteristics, themselves the products of gender-based stratification systems.

Finally, the study findings have implications for debates about defining socioeconomic position in stratification research. Consistent with feminist counter-proposals to "conventional" approaches, they suggest the need to consider the ways in which socioeconomic position is constituted from individuals' experiences in the labor market and in the family. Importantly, they also suggest that gender differences (and similarities) in the meaning of these and other social conditions and processes occupy a central position in such inquiry.

NOTES

1. Because specific conceptions of class are not the central focus of this paper, "class" is used loosely to refer to Marxist and Weberian perspectives.

2. Fitzgerald et al. (1998) suggest that mortality among PSID respondents may overstate attrition rates. Excluding those who died while in the PSID and adjusting for mortality among those lost to attrition (using national mortality rates based on age, sex, and race) increases the panel response rate, as of 1989, to 56 percent.

3. Most of the nonresponse is "family unit nonresponse" which may be the result of migration or mortality unknown to the PSID, but a residential move that cannot be successfully followed is also a common reason for attrition. Aside from a slight increase in the proportion of attritors who leave because of death and a slight decrease in those who do so because of mobility (both of which are likely a result of the increasing aging of the 1968 sample), the distribution of attrition by reason has remained fairly stable over time. Sample weights that adjust for observed variation in nonresponse are available in the PSID, but attrition occurs most frequently among those at the lower end of the socioeconomic hierarchy and those with unstable earnings, marriages, and migration histories. However, the authors of an in-depth examination of sample attrition in the PSID concluded that this nonresponse bias is moderated over time by regression-to-the-mean effects that cause initial differences to fade over time (Fitzgerald et al. 1998). In their view, attrition in the PSID has not seriously distorted its representativeness.

4. Although cause of death was not ascertained, prior research suggests that the relationship between socioeconomic position and the major causes of death is consistent
with that observed for all-cause mortality (Haan, Kaplan, and Syme 1989).

5. A total of 1,565 deaths between 1972 and 1991 (inclusive) were recorded, but only 1,461 were used in the analysis. Among the 104 individuals whose death data were not used, a range of more than two years was given as the "year" of death in 10 cases. For the remainder, vital status information was recorded after they had left the study so that up-to-date information on the covariates was unavailable.

6. National Death Index (NDI) records could not be used to verify all deaths in our study. The NDI data base was initiated in 1979, well after the beginning of our mortality followup. In addition, matching reported deaths to NDI records is complicated by the lack of social security numbers and date of birth of respondents (until 1983, age, but not date of birth, was ascertained).

7. Missing-item information on earned income varies from 2% to 6% across the study years. Standard imputation techniques, including the use of the prior year's income, result in virtually no missing values for this variable.

8. When the outcome event is a probability \( (p) \) bounded by 0 and 1, the logistic distribution or log \( (p/(1-p)) \) is often used to represent this nonlinear function.

9. When PSID began data collection in 1968, some questions, such as those concerning income, referred to the prior calendar year (i.e., 1967), while others, such as those related to disability, age, and marital status referred to the current interview year. This trend has continued in the study, resulting in differential time lags in the variables.

10. PSID weights are adjusted over time to take into consideration differential mortality and nonresponse. Differential mortality by race, sex, and age is determined by comparisons with national mortality data, while differential nonresponse is estimated as a function of prior socioeconomic characteristics, such as age, race, sex, family income, family structure, and area of residence (Institute for Social Research 1992).

11. The software package, SAS, was used to analyze the data. It computes the variance of estimates based on the assumption that sample elements are selected via simple random sampling techniques. However, the clustered nature of sample selection for the PSID renders the sample elements non-independent of one another. If this sample design feature is not taken into consideration in variance estimation, the precision of estimates is affected, generally by underestimating variance. As one of several replication techniques, the jackknife method calculates variance using information about the clustered nature of the sample design. This method was applied to the analyses by using a SAS macro program (Berglund 1997) that entails repeatedly selecting subsamples from the data and computing weighted survey estimates for each replicate. The variance is then determined from the deviations of the replicate measures from the total sample estimate (Carlson 1998).

12. Interaction terms were added to the analytic model on the basis of hierarchical ordering of terms. Hence, the interaction between gender and marital status is not of critical substantive interest but appears in the model because higher order interaction terms involving gender, marital status, and spousal education and gender, marital status, and spousal earnings are key analytic variables.

13. Specifying family income in households where married women are not engaged in paid labor is a reasonable representation of the earned incomes of male spouses.

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