
SEX DIFFERENCES IN MORBIDITY AND MORTALITY*

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Women have worse self-rated health and more hospitalization episodes than men from early adolescence to late middle age, but are less likely to die at each age. We use 14 years of data from the U.S. National Health Interview Survey to examine this paradox. Our results indicate that the difference in self-assessed health between women and men can be entirely explained by differences in the distribution of the chronic conditions they face. This is not true, however, for hospital episodes and mortality. Men with several smoking-related conditions—including cardiovascular disease and certain lung disorders—are more likely to experience hospital episodes and to die than women who suffer from the same chronic conditions, implying that men may experience more-severe forms of these conditions. While some of the difference in mortality can be explained by differences in the distribution of chronic conditions, an equally large share can be attributed to the larger adverse effects of these conditions on male mortality. The greater effects of smoking-related conditions on men's health may be due to their higher rates of smoking throughout their lives.

Research on sex differences in health in a large number of countries has brought to light an important paradox: women use more health services and report worse self-rated health than men, but women are less likely to die than same-aged men throughout life, indicating that they may, in fact, be healthier. This paradox has been the subject of earlier articles, starting with that by Nathanson (1975) and followed by numerous others (see, e.g., Idler 2003; MacIntyre, Ford, and Hunt 1999; Molarius and Janson 2002; Verbrugge 1989). In this article, we revisit this issue using more recent and larger data sets that contain information on both mortality and morbidity.

There are several possible explanations for the worse self-rated health but lower mortality of women. First, there may be sex differences in the distributions of chronic conditions, driven by biological, behavioral, or psychosocial factors (Lawlor, Ebrahim, and Davey Smith 2001; Molarius and Janson 2002; Verbrugge 1989). Women may be more likely than men to suffer from health conditions, such as arthritis or headaches, that result in poorer self-rated health but contribute relatively little to the risk of mortality, whereas men may be more likely to have conditions such as cardiovascular disease (CVD) or respiratory conditions that not only contribute to worse self-rated health but also have relatively large effects on the probability of death. Although this explanation does not address *why* women and men have different distributions of health conditions, it may account for the sex differences in self-rated health and mortality that have been observed.

Second, women may be healthier than men (as evidenced by their lower mortality), but simply *report* worse health on surveys. A commonly held view is that since women are less stoical than men, they are more likely to factor less-serious ailments into their reports of poor health (Spiers et al. 2003). A twist on the same idea is that women are more accurate reporters of health than are men: they know more about their own health,

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perhaps because of their greater use of health care, and are more willing to discuss their health and admit health problems to interviewers (Idler 2003; Verbrugge 1989). Although these ideas have been expressed in many articles, they are not uncontroversial. Some evidence that is contrary to this view was reported in a study that asked men and women open-ended questions about health problems, followed by probes for specific health conditions, and found that men provided more complete information than women in response to the open-ended questions (MacIntyre et al. 1999). Although this evidence was not paired with "objective" data on health that could be used to assess the accuracy of reports, it suggests that men are not less forthcoming than women in interviews.

Third, the "facts" underlying the paradox may be incorrect. Although the higher mortality of men is not in dispute, some researchers have questioned whether women have worse self-rated health and, more generally, higher rates of many measures of morbidity than do men (Hunt and Annandale 1999). For example, MacIntyre, Hunt, and Sweeting (1996) found evidence from two (relatively small) British samples that sex differences in reports of fair or poor health are significant only in early adulthood and that although women have higher rates of psychological distress than men, they do not have higher rates of many specific physical conditions. Some researchers have found that sex differences in self-rated health become smaller at older ages (Case and Deaton 2003), and others have found that sex differences in self-rated health disappear in old age (Arber and Cooper 1999; Leinonen, Heikkinen, and Jylha 1997).

The first explanation of the paradox—that sex differences in self-rated health and mortality are driven by differences across women and men in the distributions of chronic conditions—is consistent with women having higher rates of some illnesses than men, but lower rates of others. In fact, this explanation *requires* that sex differences in prevalence rates vary across health conditions. Much of the literature that has disputed whether women have higher rates of morbidity than men has been concerned with establishing that women do not suffer from excess levels of *all* measures of ill health—a point with which we agree and that we later provide evidence in support of. We are concerned with the more precise issue of why women have worse self-rated health but lower mortality than men. As part of our investigation of sex differences in self-rated health and mortality, we also examine sex differences in hospital episodes. Decisions about overnight hospital stays typically require the participation, if not the assent, of physicians. Measures of hospital episodes may therefore be more likely than self-rated health status to indicate the presence of a serious illness, rather than factors such as the lack of "stoicism" or the willingness of individuals to discuss their health with interviewers.¹

The work we present here is related to but distinct from a large body of literature that has investigated sex differences in the relationship between self-rated health and subsequent mortality (for reviews of studies on this topic through the mid-1990s, see Idler and Benyamini 1997; for more-recent evidence, see Deeg and Kriegsman 2003, Idler 2003). These studies have examined whether the association between self-rated health and mortality, adjusted for sociodemographic characteristics and chronic health conditions, varies across men and women. A general (although not universal) finding has been that with few controls included, the association between self-rated health and subsequent mortality is larger for men than for women. This finding has been taken as evidence that women have different health-reporting behaviors or different health knowledge than men. However, in much of this literature, as more controls for health conditions are added, the associations between self-rated health and mortality decline for both women and men, and the gap between men and women becomes small and often insignificant. This finding suggests

1. Sex differences in hospitalization episodes could partly reflect women's greater willingness to be hospitalized or to seek medical care that results in hospital stays. We discuss this possibility in more detail later.

that differences in the distributions of health conditions across men and women may play a large part in resolving the paradox of higher morbidity but lower mortality among women—something we explore here.

Our analyses are based on 14 years of U.S. data from the National Health Interview Survey (NHIS 1986–1994, 1997–2001) and its associated Multiple Cause of Death file, which provides information on the deaths of those who were interviewed between 1986 and 1994. The surveys are ideal for our purpose because they contain information on self-rated health, hospital episodes, and mortality, as well as detailed information on chronic health conditions that may contribute to both self-rated health and mortality. In addition, the survey covers a large number of men and women: our analyses of self-rated health and hospitalization use a sample of 147,996 men and women who were surveyed between 1997 and 2001, and our analyses of mortality use a sample of 237,140 men and women aged 45–84 who were surveyed from 1986 to 1994. These large samples make it possible for us to obtain precise estimates of the sex differences in health measures at each age and of the effects of (sometimes rare) chronic health conditions on self-rated health and mortality.

In the next section, we discuss the data and show that they are consistent with the basic facts of the puzzle: women have worse self-rated health and more hospitalization episodes than men from early adolescence to late middle age, but they are less likely to die at each age. We then examine the validity of the explanations offered for this paradox. Our results indicate that the difference in self-assessed health between women and men can be entirely explained by differences in the distribution of conditions. Although women have, on average, worse self-rated health than men, women and men with the same sets of chronic conditions have the same self-rated health. The results for hospital episodes are somewhat different. While the effect of poor health on hospital episodes is the same for men and women, men with respiratory cancer, CVD, and bronchitis are more likely to experience hospital episodes than women with the same conditions, implying that men may experience more-severe forms of these conditions. The same is true for mortality. Although the effects of many chronic conditions on the probability of death are the same for women and men, men who report having CVD and certain lung disorders are significantly more likely to die than women with these conditions. While some of the sex difference in mortality can be explained by differences in the distribution of chronic conditions, an equally large share can be attributed to the larger adverse effects of these conditions on male mortality.

Our results move us some distance toward understanding the paradox between self-assessed health and mortality. That men and women with the same health conditions report the same self-rated health status and that poor health is equally predictive of hospitalization episodes for men and women cast doubt on the idea that women and men use different standards for assessing self-rated health. However, our mortality results cannot be explained solely by differences in the distribution of chronic conditions. Men with smoking-related conditions are significantly more likely to die within two years than are women with the same conditions. These men will, on average, have had longer exposure to smoking in their lives, and their reports of, say, emphysema may indicate more advanced cases than do women's reports.

Our findings and their interpretation are important for several reasons. Self-reported health status is a tool that is often used to assess well-being. If women and men take different aspects of health into account, or weight them differently in their self assessments, it is important to understand why and how they do so. In addition, if historically higher rates of smoking are responsible for the higher mortality rates that men face in middle age, then we anticipate that the gap in age-adjusted mortality rates will close, given the changes that have been observed in women's and men's smoking patterns over the past century.

DATA AND PRELIMINARY EVIDENCE

Data

The data for this study are drawn from the NHIS from 1986 to 2001, and from the associated NHIS Multiple Cause of Death Public Use Data File, which contains information on the deaths (as of 1997) of individuals who were surveyed between 1986 and 1994. The NHIS is a cross-sectional household interview survey that covers the civilian noninstitutionalized population of the United States. The survey collects information on self-rated health, chronic health conditions, the use of health care services, and socio-demographic characteristics.

Between 1986 and 1996, health information was collected for each member of the sampled households. Although all households were administered the same basic questionnaire, each household was randomly assigned to one of six "Condition Lists," and information was collected only on household members' experience with the chronic conditions that were included in the assigned list. The structure of the survey changed in important ways in 1997. While basic health information continued to be collected for all household members, information on chronic conditions among adults was collected only from a single "sample adult" in each household, who was asked about a full range of chronic conditions.

Because of the redesign and the timing of the Multiple Cause of Death File, we rely mainly on two distinct samples to study self-rated health and mortality. The 1997–2001 sample, consisting of 147,996 "sample adults" (men and women) aged 18–84, is used mainly for our analyses of self-rated health. We remove women who were either pregnant or had a child aged 1 or younger, to focus on non-pregnancy-related health and hospitalizations. The fact that we have a complete set of information on chronic conditions on all sample adults is an advantage, since it permits us to deal more easily with comorbidities across chronic conditions. We use the 1986–1994 sample primarily for our analyses of mortality. That sample includes men and women whose vital status can be identified from the Multiple Cause of Death File. Because there are so few deaths among younger adults, we restrict our sample to 237,140 men and women aged 45–84 whose vital status is known. Details on the definitions of the variables are included in Appendix A.

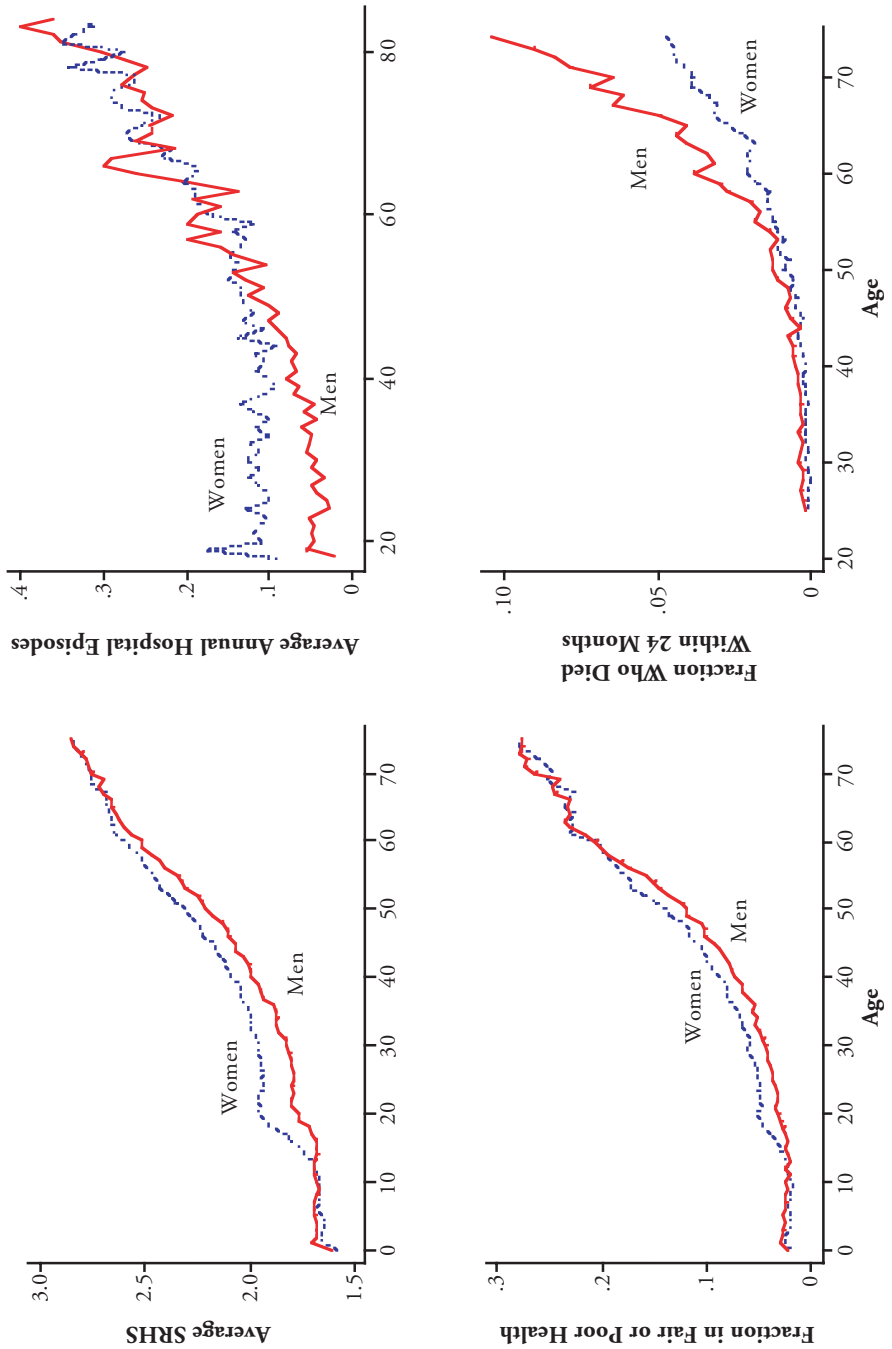
Sex Differences in Self-Rated Health and Mortality

The NHIS data can be used to illustrate the paradox that we discussed in the previous section. The two left-hand panels of Figure 1 present data on self-rated health status (SRHS), which was coded on a 5-point scale (1 = excellent, 2 = very good, 3 = good, 4 = fair, and 5 = poor). The figure, which is based on data on all household members from the combined 1986–2001 surveys, shows the average SRHS and the proportions of men and women who rated themselves in fair or poor health, by exact age.² These measures yield similar patterns of changes in health status with age. Boys have somewhat worse health status than girls prior to adolescence, but the health status of girls becomes worse than that of boys at about age 14. The sex gap in self-rated health status is greatest at age 20 and then slowly declines. By age 65 (using the mean SRHS) or age 60 (using the indicator of fair or poor health), the female disadvantage in self-rated health status has vanished.

The top right panel shows the average number of hospitalization episodes over the previous 12 months, by age, for women and men. Between ages 20 and 60, women's number of hospitalizations is roughly constant, at 0.10 episodes per year. Only after age 60 does the average number of episodes rise with age for women. In contrast, men's

2. These results are based on data for 1,602,650 individuals. Sample weights are used in the calculations. The health of those aged 16 and younger was rated by adult respondents, but children aged 17 had the option of reporting for themselves.

Figure 1. Self-Rated Health, Hospital Episodes, and Mortality for Men and Women: NHIS



reports of hospitalizations rise with age from age 20. The pattern here mirrors what we observed for self-assessed health: women's excess hospitalizations are greatest at age 20, but men's and women's reported hospitalizations equalize by late middle age.

The bottom right panel of the figure presents the proportion of respondents in the 1986–1994 surveys (aged 25–74 at the time of the survey) who died within 24 months of their interviews.³ It shows the well-known pattern of excess mortality among men at all ages. However, the scale of the figure, with low mortality at the youngest ages, makes it difficult to discern how the ratio of male-to-female mortality changes with age. A graph of the ratio of male-to-female mortality (not shown) indicates that the “excess” of male mortality is the greatest at the youngest ages and is fairly flat after age 40, increasing slightly at the oldest ages.

Although women report worse health than men at all but the oldest ages, they are not uniformly more likely to report that they have all chronic health conditions. Figure 2 presents the prevalence rates for a selection of chronic health conditions for women and men by exact age. These chronic conditions were chosen to illustrate the diversity of patterns that appear in the data. Several chronic conditions, including frequent headaches, arthritis, and depression, are more prevalent for women at all ages. For others, such as asthma, there is a higher prevalence among women than among men in middle age but not at the oldest ages. Yet others, such as reproductive cancers and CVD, are more prevalent among women at younger ages but become more prevalent among men at older ages. Diabetes is equally prevalent for men and women at all ages. Emphysema and, to some extent, hypertension are equally prevalent among younger men and women but are more prevalent among men older than age 60. These varied patterns confirm the view that women are not uniformly more likely than men to suffer from all types of ailments. However, they also suggest that women are more likely than men to suffer from conditions, such as arthritis and frequent headaches, that are not life threatening but could lead to poor self-rated health and that men are more likely, at least at older ages when most deaths occur, to suffer from diseases, such as CVD and emphysema, that are risk factors for mortality.

CHRONIC CONDITIONS, SELF-RATED HEALTH, AND MORTALITY

Self-Rated Health

Methods. To assess the routes to women's poorer health, we examine the relationship between reports of chronic conditions and self-rated health using the 1997–2001 surveys. Here, we present Oaxaca-like decompositions of the sex difference in health status into three components: a “prevalence effect,” which is due to sex differences in the distributions of chronic conditions; a “severity effect,” which is due to sex differences in the impact of these conditions; and a remainder, which is due to differences in a set of other observable characteristics (age, education, and race) and to unobserved factors. The measure of self-rated health is a binary indicator for whether self-rated health is fair or poor, which we hereafter refer to as “poor” health.

Our analysis uses the 18 chronic conditions listed in Table 1. We rule out conditions that are benign enough to pose little health hazard (e.g., hay fever), and restrict our attention to conditions for which a consistent definition was used in all the years. These 18 conditions fall into six broad categories: pain (headache; arthritis; and “other,” including neck, back, and joint pain that is not due to arthritis), respiratory conditions (bronchitis, emphysema, asthma, and lung problems), circulatory conditions (CVD, diabetes, hypertension, and circulatory problems), cancers (skin, stomach, reproductive, and respiratory), chronic vision and hearing problems, and depression. The list of conditions is made up of

3. This figure is based on 586,703 respondents of all ages whose vital status could be determined.

Figure 2. Prevalence of Selected Conditions, by Age

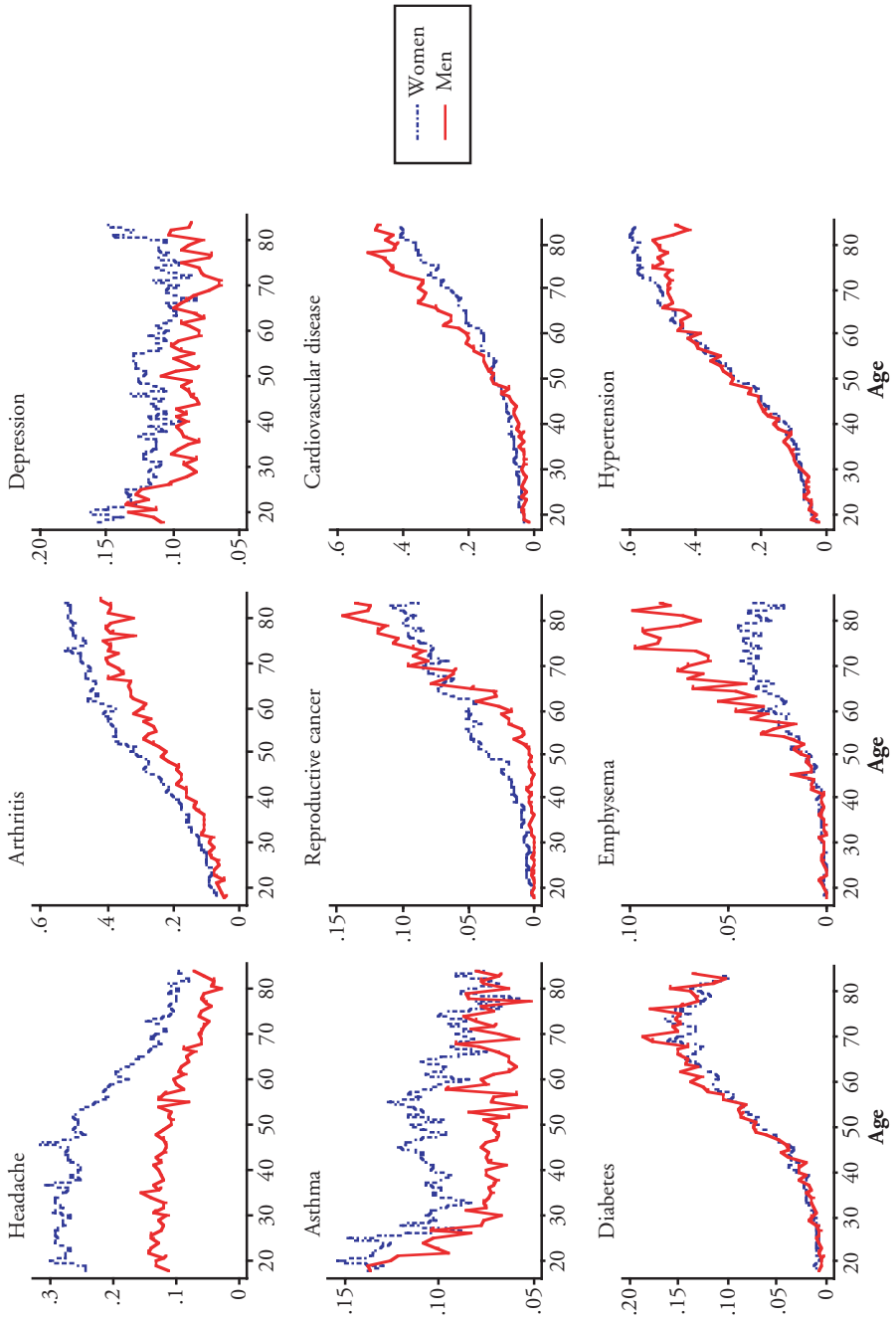


Table 1. Prevalence Rates of Chronic Conditions (percentages)

Condition	Prevalence, Women	Prevalence, Men	Excess Prevalence in Women (percentage points)
Headache	24.0	11.4	13.2*
Other Pain	39.5	35.6	3.7*
Arthritis	28.0	19.2	7.2*
Bronchitis	6.7	3.2	3.3*
Emphysema	1.4	2.0	-0.8*
Lung Problems	2.0	1.4	0.5*
Asthma	10.8	7.8	3.1*
Diabetes	6.6	6.1	-0.3*
Circulatory Problems	0.5	0.4	0.05
Cardiovascular Disease	13.3	12.9	-1.1*
Hypertension	26.5	23.3	0.5*
Skin Cancer	1.9	2.3	-0.5*
Stomach Cancer	0.1	0.1	-0.01
Reproductive Cancer	3.3	1.8	1.2*
Respiratory Cancer	0.2	0.4	-0.2*
Vision Problems	11.2	8.3	2.2*
Hearing Loss	3.7	6.0	-3.0*
Depression	12.8	9.7	2.9*
Number of Observations	81,704	66,292	147,996

Notes: Excess prevalence coefficients are the coefficients on an indicator that the respondent is female in OLS regressions for each condition. Each regression also includes a complete set of age, survey year, and race indicators, and a variable for completed education. Because these prevalence rates are adjusted by age, race, and education, they do not correspond to crude differences in male-female prevalence rates.

*The difference in the rate for men and women is significant at the 5% level.

both diseases (e.g., emphysema) and illnesses that may be symptoms of diseases (e.g., vision problems may be due to diseases of the eye or may be a by-product of diabetes).

We model the probability of being in poor health as a linear function of indicators for the presence of a set of N chronic conditions (denoted as C_i , $i = 1 \dots N$); indicator variables for age, race, and survey year; and a control for education (X). The coefficients in the model are assumed to differ across women (W) and men (M):

$$P(\mathbf{H}) = \sum_i^N \beta_i^j C_i + \mathbf{X}\boldsymbol{\gamma}^j + \boldsymbol{\varepsilon}, \quad j = W, M. \quad (1)$$

We estimate (1) separately for women and men, using linear probability models.⁴ Estimates of β^W and β^M provide information on sex differences in how reports of chronic conditions map into reports of poor health. In addition, the parameter estimates, together with information on the prevalence of the chronic conditions, can be used to construct

4. We use ordinary least squares (OLS) to estimate Eq. (1) because doing so results in an exact linear decomposition of sex differences into three additive components. However, marginal effects from estimates of probit models are extremely similar to the OLS estimates we present.

severity and prevalence effects. The severity effect is measured as the sum of differences in the coefficients for women's and men's chronic conditions, multiplied by the average prevalence rate of each condition over both men and women (denoted as \bar{C}_i):

$$\text{severity effect} = \sum_i (\beta_i^W - \beta_i^M) \bar{C}_i.$$

The prevalence effect is measured as the sum over conditions of differences in prevalence rates between women and men, multiplied by the condition's β (averaged between the β estimated for men and women):

$$\text{prevalence effect} = \sum_i (\bar{C}_i^W - \bar{C}_i^M) \bar{\beta}_i.$$

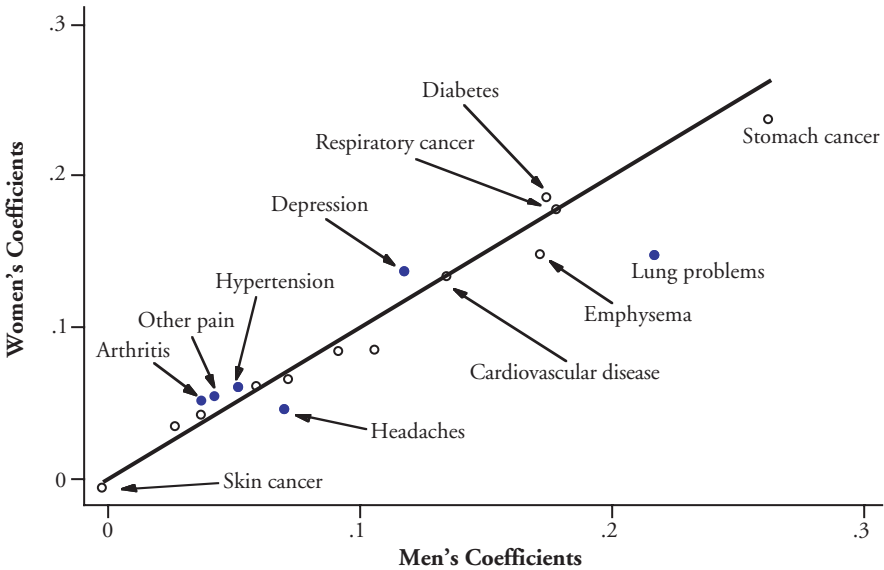
The residual difference in women's and men's self-assessed health—the portion that is unexplained by the severity and prevalence effects just presented—combines differences that are produced by sex differences in average values of X (age, education, and race), sex differences in the effects of these variables on health (e.g., in γ), and unexplained differences.

We estimate two variants of Eq. (1) that differ in their measures of chronic conditions. In the first variant, each of the N indicators of chronic conditions measures the presence or absence of one of the 18 ailments in our list (e.g., arthritis or asthma). This specification is parsimonious, but does not account for the fact that health status may be influenced by the specific combination of conditions an individual has. In other words, the probability of poor health, given two chronic conditions, may not be the simple sum of the probability of poor health for each chronic condition observed individually. In our data, 45% of the respondents suffer from more than one chronic condition. The second variant allows for a large number of combinations of chronic conditions. In theory, we could incorporate the presence of multiple conditions by redefining our chronic conditions to measure each unique combination of ailments. In practice, it is not possible to allow for all possible combinations of the 18 ailments we measure. There are, for example, 153 different ways in which a person could suffer from exactly 2 of the 18 ailments and 816 different ways in which a person could suffer from exactly 3 ailments. Instead, we capture the first-order interactions between conditions by defining “conditions” as single chronic ailments and all possible combinations of 2 chronic ailments. Individuals who suffer from CVD and emphysema have a “CVD-emphysema” condition. To avoid double counting, we do not assign these individuals to the “CVD only” or “emphysema only” conditions. We thus define 171 new conditions: 18 singletons plus 153 pairs of ailments (so that $N = 171$). For the 73% of the individuals who reported 2 or fewer ailments, the conditions we identify provide a complete accounting of the combinations of ailments that they have. For the remainder, the estimation yields an approximation of the effects of their combinations of conditions on health.

Results. Figure 3 graphs estimates of β^W (on the y -axis) against estimates of β^M (on the x -axis) from the first variant of Eq. (1), in which the list of conditions consists of the 18 conditions and no combinations of conditions. For both men and women, the condition that has the largest effect on self-assessed health is stomach cancer, which is associated with an increased probability of reporting poor health of roughly 25 percentage points. Respiratory cancer and diabetes are each associated with an 18-percentage-point increase in the probability of reporting poor health, for both men and women; depression and CVD, with a 15-percentage-point increase; and other pain, with a 5-percentage-point increase. Only one condition—skin cancer—is not significantly associated with poor health; the coefficient for this condition is small and negative for both men and women.

The solid line in Figure 3 marks the 45-degree line: if the increase in the probability of reporting poor health in the presence of a given chronic condition were equal for men

Figure 3. The Impact of Chronic Conditions on the Probability of Reporting Fair or Poor Health



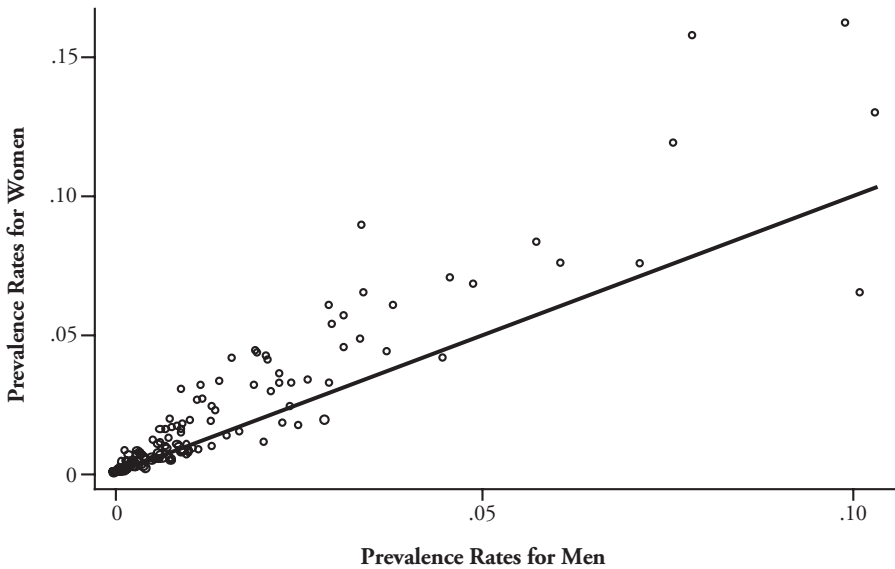
Note: Coefficients that are significantly different for men and women are represented by solid circles.

and women, then the regression coefficients for the two sexes would fall on this line. Figure 3 shows a remarkable similarity between the sexes in the association between poor health and the presence of each chronic ailment. Tests for differences between men and women in the 18 ailment coefficients are significant for only 6: headaches, other pain, arthritis, hypertension, depression, and lung problems. Of these 6, only the difference associated with lung problems is large in magnitude, where the probability of poor health is 7.0 percentage points higher for men than for women. In what follows, lung disorders and, more generally, chronic illnesses that are associated with smoking will be seen to affect hospitalizations and mortality more for men than for women.

Estimates of the second variant of Eq. (1), which include 18 conditions and 153 combinations of conditions, yield similar results. Although some of the individual coefficients on chronic conditions are imprecisely estimated (because some combinations of conditions are rare), the coefficients for women and men are again remarkably similar. The inclusion of indicators of combinations improves the fit of the model, and the hypothesis that there are no interaction effects in conditions is rejected.

Although the associations between chronic conditions and poor health are similar for men and for women, there are marked sex differences in the prevalence of conditions. Table 1 presents the prevalence rates for chronic conditions and the differences in the prevalence rates between women and men for the 18 ailments we use in our analysis. The last column shows the age- and race-adjusted excess percentage-point prevalence among women for each ailment, above that expected for men. These prevalences were estimated by regressing each ailment on a complete set of age, race, and survey-year indicators; a measure of completed education; and an indicator that the respondent is female, for 147,996 individuals who were surveyed between 1997 and 2001. Women have significantly higher

Figure 4. The Prevalence of 171 Chronic Conditions and Condition Interactions for Men and Women Aged 18–64: NHIS, 1997–2001



rates of pain (headache, other pain, and arthritis) and some respiratory conditions (bronchitis, asthma, and lung problems other than cancer) and are significantly more likely to suffer from reproductive cancers, hypertension, vision problems, and depression than men. Men are significantly more likely than women to suffer from smoking-related ailments (emphysema and respiratory cancer), some circulatory problems (CVD and diabetes), and hearing loss. There are also differences between men and women in the prevalence of combinations of conditions. Men are significantly more likely than women to have two-ailment conditions that involve emphysema, respiratory cancer, diabetes, and CVD, whereas women are significantly more likely than men to have two-ailment conditions that involve pain, asthma, bronchitis, vision problems, and depression. The prevalence rates for the set of 171 chronic conditions and combinations of conditions are presented in Figure 4, where it is clear that the difference in the prevalence rates for women and men are larger for conditions that women are more likely to report.

Estimates of the prevalence and severity effects are presented in Table 2. The first column shows the results based on regressions that include 18 chronic conditions and no combinations of conditions. The second column shows the results based on regressions that included indicators of 171 chronic conditions and combinations of conditions. On average, 14.3% of the women aged 18–84 reported fair or poor health. The likelihood of a woman reporting poor health is 2.5 percentage points greater than it is for a man. With or without the combinations of chronic conditions, the difference between women's and men's health is fully explained by differences in the prevalence of the conditions they have. From our analysis that includes combinations of conditions (column 2), we predict that women are 2.3 percentage points more likely to report poor health on the basis of the conditions they report. There is little role for the severity effect: men are 0.13 percentage points more

Table 2. The Decomposition of Health Status

	Fraction of Respondents Who Reported Fair or Poor Health	
	Allowing for 18 Conditions	Allowing for 18 Conditions + 153 Condition Combinations
Fraction Who Reported Fair or Poor Health		
Women	0.1428	0.1428
Men	0.1181	0.1181
Difference (women – men)	0.0247	0.0247
Decomposition of Difference		
Severity effect	0.0065	-0.0013
Prevalence effect	0.0261	0.0233
Residual difference	-0.0079	0.0028
<i>F</i> Test for the Joint Significance of 153 Condition Interaction Terms (<i>p</i> value) for Women's Regression		8.07 (0.0000)

likely to report poor health than are women with the same chronic conditions. Although the inclusion of combinations of conditions improved the fit of Eq. (1), excluding these combinations (column 1) makes little difference to the health-status decomposition.

Hospitalizations

An alternative measure of morbidity is provided by the number of in-patient hospitalization days and episodes that men and women report having had in the past 12 months. As is shown in Figure 1, women experience more hospitalizations than do men, especially at younger ages. Hospitalizations may provide a more objective measure of morbidity than do self-reports of health. Hospital admissions and the length of hospital stays are controlled by health care providers and insurers, watchdogs who presumably base their decisions on the medical needs and the severity of patients' illnesses. However, hospitalization could also contain a subjective element. It is possible that, when the nature of their illnesses is held fixed, women are more likely than men to seek medical care that results in hospitalization or are more willing than men to be hospitalized—or that health care providers are, for some reason, more likely to recommend hospitalization for women than for equally sick men. Although whether hospitalizations represent a more objective measure of morbidity than do self-reports of health status remains an open question, hospitalizations certainly represent an alternative measure of morbidity.

Before we describe the results of our decomposition analyses, we examine the relationship between poor health and hospitalizations. If women are more likely than men to factor minor ailments into their reports of poor health and if hospital episodes are good indicators of serious illness, then we would expect women's hospitalizations to be less responsive to women's reports of poor health than men's. Instead, our results show equally large responses between hospitalization episodes and reports of poor health for men and women. The results of regressions of the number of reported hospitalization episodes on an indicator of reporting of fair or poor health, with a complete set of indicators for age, race, and survey year and a measure of educational attainment reveal the following:

Women: Number of episodes = 0.341 Fair or Poor Health (0.006)

Men: Number of episodes = 0.338 Fair or Poor Health (0.007)

For both women and men, a report of fair or poor health translates, on average, into one third of an extra hospitalization episode per year.

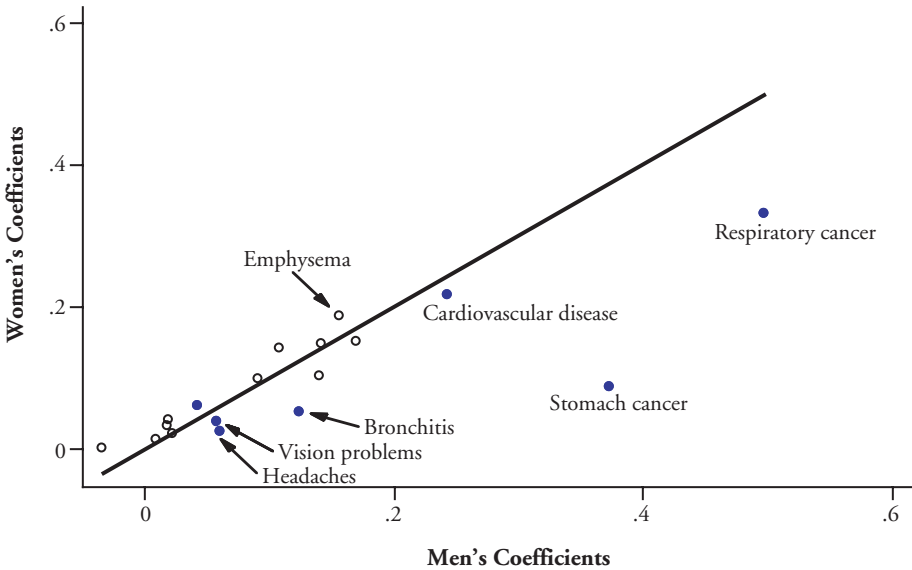
Using the same methods presented for health status, we decompose hospitalization days and episodes into severity, prevalence, and residual effects. We present these decompositions in Table 3, where the first three columns indicate the results for hospitalization episodes and the second three indicate the results for hospitalization days. For both episodes and days, we present the results for our 18 chronic conditions without condition combinations (columns 1 and 4) and with the combinations (columns 2, 3, 5, and 6). Women, on average, have 0.034 more hospitalization episodes and 0.052 more hospital days per year than do men. Consistent with our findings for self-reported health status, the difference in the distribution of chronic conditions goes some way toward explaining women’s greater hospitalization. The prevalence effect for hospitalization episodes is 0.0176, based on our analysis containing the 171 conditions and condition combinations. This is half the total difference we observe in hospitalization episodes between men and women. The prevalence effect for hospital days is 0.0941, almost twice as large as the total gap between women’s and men’s hospitalization days. On the basis of the distribution of conditions alone, we would expect women to spend even more days in the hospital, relative to men, than we observe in the data.

These prevalence effects are offset by differences in the impact of chronic conditions on hospitalizations for men and women. For both episodes and days, when men and women report the same chronic conditions, men have more episodes and more hospital days than do women. The conditions that are responsible for this severity effect can be seen in Figure 5, which presents coefficients from the regressions of men’s and women’s hospitalization episodes on 18 chronic conditions. Men who have respiratory cancer, stomach cancer, bronchitis, and CVD have a significantly greater number of hospital episodes than do women who have the same conditions.

Table 3. The Decomposition of Hospitalization Episodes

	Hospitalization Episodes			Hospitalization Days		
	18 Conditions, All Ages	18 Conditions + 153 Condition Interactions		18 Conditions, All Ages	18 Conditions + 153 Condition Interactions	
		All Ages	Ages 45–84		All Ages	Ages 45–84
Averages						
Women	0.1502	0.1502	0.1870	0.7286	0.7286	1.0190
Men	0.1160	0.1160	0.1853	0.6769	0.6769	1.0940
Difference	0.0343	0.0343	0.0016	0.0517	0.0517	–0.0749
Decomposition of Difference						
Severity effect	–0.0133	–0.0082	–0.0030	–0.1170	–0.0861	–0.0800
Prevalence effect	0.0196	0.0176	0.0089	0.0980	0.0941	0.0268
Residual difference	0.0252	0.0241	–0.0045	0.0642	0.0362	–0.0281

Figure 5. The Impact of Chronic Conditions on Hospital Episodes in the Past 12 Months



Note: Coefficients that are significantly different for men and women are shown as solid circles.

Mortality

Methods. The previous two sections provided evidence that the worse self-rated health of women can be entirely explained by sex differences in the distribution of chronic health conditions, but that men are more likely than women with the same health conditions to have hospitalizations. This section turns to mortality. We examine whether men and women with the same health conditions are equally likely to die in the two years following the survey or whether men have excess mortality, given their health conditions. We also decompose sex differences in mortality into prevalence and severity effects, as was done for self-rated health and hospitalizations.

As in the previous sections, we start by specifying mortality as a function of health conditions (C_i) and other sociodemographic factors (X):

$$P(D) = \sum_i \beta_i' C_i + X\gamma^j + \epsilon = C\beta^j + X\gamma^j + \epsilon, \quad j = W, M, \tag{2}$$

where D is an indicator that the respondent died within 24 months of the survey. These equations are estimated separately for women and men. Because there are few deaths among younger adults, the samples include only those aged 45–84. Our primary focus is on whether the effects of chronic conditions on mortality differ for women and men.

Although the framework is identical to that described in the previous sections, data issues require us to use different and somewhat more complex estimation methods. The NHIS mortality information is available only for those who were surveyed before the 1997 redesign. As we discussed in the Data section, the NHIS followed different procedures for collecting information on health conditions before 1997: instead of asking a

subset of adults about all health conditions, the survey asked all adults about a subset of health conditions on the list of conditions to which they were randomly assigned. As a result, we have only incomplete information on the variables in \mathbf{C} for any one individual. Although this method of collecting information allows for the calculation of accurate prevalence rates with a minimal burden on the respondents, it complicates analyses that require information on comorbidities. Consider, for example, the strategy of estimating Eq. (2) one list of conditions at a time, so that we regress the indicator of death on the conditions that are included in Condition List 1 using the sample that was assigned to this list and repeat this procedure for each list of conditions. The estimates of the effects of the conditions on the probability of mortality (i.e., the estimates of β) would be biased, since they will reflect the effects of conditions that are unmeasured but are correlated with the conditions that are measured. For example, if those who have heart disease also have an elevated risk of emphysema (perhaps because smoking increases the risk for both diseases), the estimates of the effects of both heart disease and emphysema, which appear in different condition lists, would likely be biased upward.

The bias that is due to incomplete information on chronic conditions can be corrected with supplementary information on the covariance of each pair of chronic diseases that appear in the different lists of conditions. Our strategy is to use information on these covariances from the 1997–2001 NHIS surveys to correct for bias that is due to missing information on comorbidities.⁵ To see how these biases can be corrected, consider a simplified version of Eq. (2) in which the controls for the sociodemographic variables have been suppressed:

$$\mathbf{D} = \mathbf{C}_1\beta_1 + \mathbf{C}_2\beta_2 + \dots + \mathbf{C}_M\beta_M + \boldsymbol{\varepsilon} = \mathbf{C}\beta + \boldsymbol{\varepsilon}, \quad (3)$$

where \mathbf{C}_j represents an $N \times k_j$ matrix of indicators for whether individuals have chronic conditions included in condition list j , and there are a total of M condition lists. (In practice, we include controls for sociodemographic variables. The methods described next are easily extended to the case in which other control variables are included.) Note that the 1986–1994 data, which we refer to as the “censored” sample, contain information on \mathbf{D} but incomplete information on \mathbf{C} . The 1997–2001 data, which we refer to as the “supplemental” sample, has complete information on \mathbf{C} but no information on \mathbf{D} .

The equation for bias correction can be obtained by first defining a matrix \mathbf{Z} , which is equal to \mathbf{C} but with unobserved values of conditions set to 0, and considering the OLS estimates of β that result when \mathbf{Z} is used in place of \mathbf{C} :

$$\tilde{\beta} = (\mathbf{Z}'\mathbf{Z})^{-1}(\mathbf{Z}'\mathbf{C})\beta + (\mathbf{Z}'\mathbf{Z})^{-1}\mathbf{Z}'\boldsymbol{\varepsilon} \quad (4)$$

and

$$\text{plim } \tilde{\beta} = \Sigma_{ZZ}^{-1}\Sigma_{ZC}\beta, \quad (5)$$

where Σ_{ZZ} is the variance-covariance matrix of the censored data matrix \mathbf{Z} and Σ_{ZC} is the covariance matrix of \mathbf{Z} and \mathbf{C} . Eq. (5) indicates that $\tilde{\beta}$ is an inconsistent estimator of β and suggests the following corrected estimator (denoted $\hat{\beta}$):

$$\hat{\beta} = \hat{\Sigma}_{ZC}^{-1}\hat{\Sigma}_{ZZ}\tilde{\beta}, \quad (6)$$

5. The strategy of using supplemental data in this way has been used to correct for biases in the estimates of the effects of lead on IQ (Marais and Wecker 1998).

where $\hat{\Sigma}_{ZZ}$ and $\hat{\Sigma}_{ZC}$ are consistent estimators of the covariance matrices. The former can be obtained from the censored sample. The matrix Σ_{ZC} contains cross-products for pairs of conditions that are from the same lists of conditions (and so can be estimated from the censored sample alone) and for pairs of conditions that appear in different lists (and so must be estimated from the supplemental sample). Details on how these covariance matrices were estimated, and on estimates of the standard errors for β , are in Appendix B.

Although this estimation method delivers consistent estimates of β for women and men, the lack of complete information on all conditions is restrictive. We cannot, for example, estimate models similar to those shown for self-rated health and hospitalizations, in which combinations of conditions are included. We also cannot use nonlinear functional forms for Eq. (3)—for example, logit or probit models—since the bias-correction formulas are correct only for the linear model. Finally, we must restrict the list of health conditions to those that were measured in similar ways in the earlier censored sample and the later supplemental sample. We examine 14 chronic conditions: arthritis and skin cancer (from List 1); digestive cancer (from List 3); diabetes, frequent headaches, and reproductive cancer (from List 4); CVD, hypertension, and other circulation problems (from List 5); and bronchitis, asthma, emphysema, other (noncancerous) lung disorders, and respiratory cancer (from List 6).

Results. Table 4 presents the results from two sets of estimates of the effects of chronic conditions on the probability of death within 24 months of the survey. The first two columns show estimates of simple variants of Eq. (2), in which the indicator for death is regressed on only *one* of the indicators of chronic conditions, plus a set of indicators for age and race and a control for years of education. Thus, each cell of each row shows the marginal effects of the condition on the probability of death for women (column 1) and men (column 2), controlling for no other conditions. The third column provides the p value from a test of whether the coefficients for women and men are identical. These simple results are shown so that they can be contrasted with those in the last three columns, which contain results from the complete model that includes the bias correction we discussed earlier.

The results with and without bias correction are generally consistent with those shown for self-rated health and hospitalizations. The chronic conditions that lead to the largest increases in the probability of death—CVD, emphysema, other lung problems, respiratory cancer, digestive and reproductive cancers, and diabetes—also elevate reports of poor health and hospitalizations. Conditions that have smaller or even no effects on mortality, such as headaches or hypertension, have smaller (although significant) adverse effects on self-rated health.

Estimates of the marginal effects of conditions on mortality are generally smaller when the complete set of conditions is included and the bias correction is used. For example, CVD is estimated to increase the probability of death by 2.9 percentage points for women and 4.6 percentage points for men when no conditions other than CVD are included. In the bias-corrected model, these effects decline to 2.1 percentage points (for women) and 3.6 percentage points (for men). These decreases reflect the fact that individuals often have multiple conditions, which was illustrated in the results in the previous section.

Correcting the bias tends to reduce the men's coefficients by more than the women's coefficients. For example, without adjustment for comorbidities, the difference in the effects of CVD on mortality is -1.7 percentage points (2.9 for women minus 4.6 for men). The adjustment for comorbidities reduces this difference to -1.5 (2.1 for women minus 3.6 for men). The pattern of greater reductions in men's than in women's coefficients is evident for 10 of the 14 conditions. This finding is consistent with the evidence on patterns of comorbidities discussed in the Data section, which noted that men are

Table 4. Marginal Effects of Conditions on Two-Year Mortality

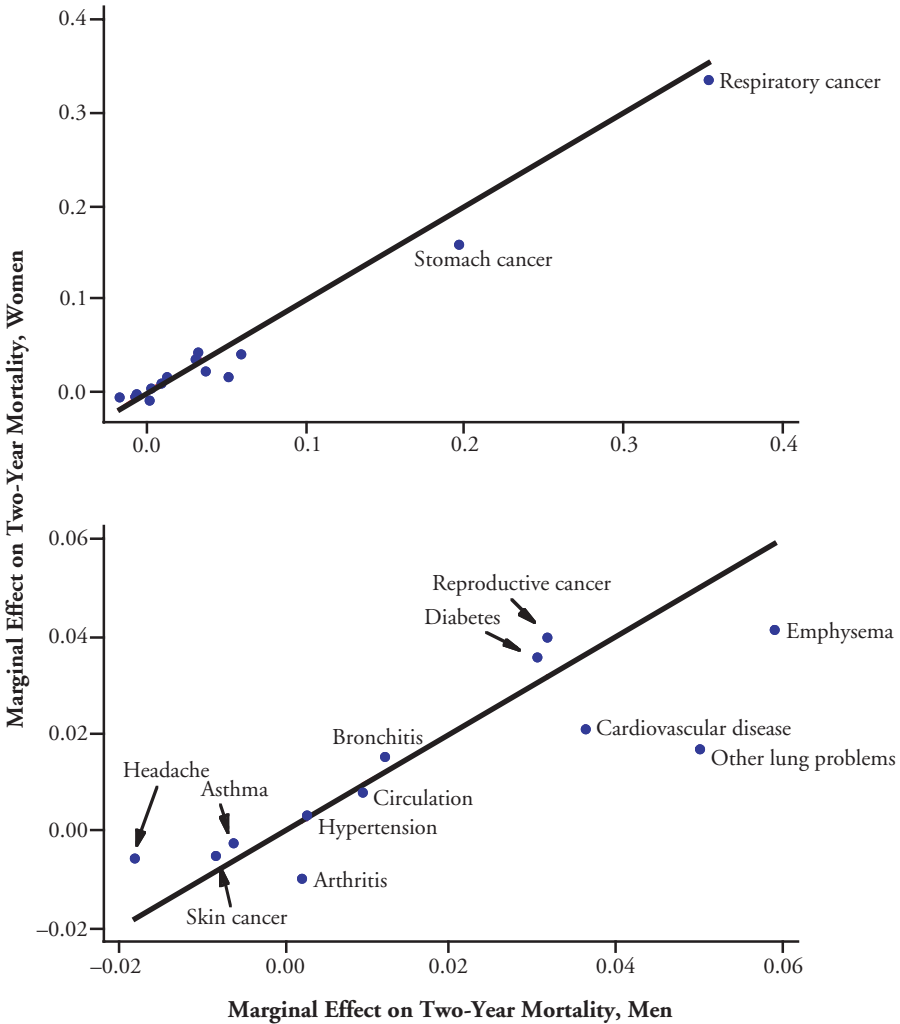
Condition	Unadjusted for Comorbidities			Adjusted for Comorbidities		
	Women	Men	<i>p</i> Value: Women = Men	Women	Men	<i>p</i> Value: Women = Men
Arthritis	-0.0032 (0.0023)	0.0067 (0.0028)	0.006	-0.0099 (0.0029)	0.0018 (0.0039)	0.016
Skin Cancer	0.0006 (0.0084)	-0.0055 (0.0069)	0.577	-0.0050 (0.0077)	-0.0083 (0.0079)	0.762
Digestive Cancer	0.1630 (0.0189)	0.1940 (0.0199)	0.259	0.1571 (0.0169)	0.1973 (0.0219)	0.146
Diabetes	0.0401 (0.0042)	0.0415 (0.0047)	0.826	0.0355 (0.0043)	0.0305 (0.0058)	0.489
Frequent Headache	-0.0015 (0.0036)	-0.0037 (0.0057)	0.746	-0.0054 (0.0037)	-0.0180 (0.0068)	0.100
Reproductive Cancer	0.0463 (0.0082)	0.0537 (0.0131)	0.633	0.0395 (0.0074)	0.0317 (0.0146)	0.632
Cardiovascular Disease	0.0289 (0.0029)	0.0459 (0.0031)	0.000	0.0212 (0.0031)	0.0362 (0.0041)	0.004
Hypertension	0.0104 (0.0025)	0.0141 (0.0028)	0.316	0.0030 (0.0028)	0.0024 (0.0037)	0.908
Circulatory Problems	0.0128 (0.0039)	0.0166 (0.0044)	0.527	0.0081 (0.0035)	0.0094 (0.0050)	0.840
Bronchitis	0.0212 (0.0043)	0.0299 (0.0059)	0.234	0.0153 (0.0041)	0.0121 (0.0068)	0.684
Asthma	0.0066 (0.0051)	0.0130 (0.0067)	0.449	-0.0026 (0.0048)	-0.0064 (0.0076)	0.666
Emphysema	0.0605 (0.0088)	0.0752 (0.0069)	0.189	0.0415 (0.0081)	0.0590 (0.0080)	0.122
Other Lung Disorder	0.0321 (0.0092)	0.0710 (0.0090)	0.003	0.0171 (0.0084)	0.0502 (0.0102)	0.012
Respiratory Cancer	0.3469 (0.0264)	0.3683 (0.0194)	0.513	0.3342 (0.0236)	0.3545 (0.0214)	0.524

Notes: The results shown in the first three columns are from linear regressions of an indicator that the respondent died within two years of the survey on an indicator that the respondent had the condition listed in the row, plus a set of age indicators, indicators for race (black or other nonwhite), and years of education. Separate regressions were estimated for each condition. Each cell shows the regression coefficient and standard error for the condition listed in the row. The third column is the *p* value for a *t* test of the hypothesis that the effects of each condition on male and female mortality are identical. The results in the last three columns are from regressions of an indicator that the respondent died within two years of the survey on a set of indicators for whether the respondent had each of the conditions, plus a set of age and race indicators and a control for years of education. These estimates have been corrected to account for the fact that not all conditions are observed for each individual, using the procedure described in the text and in Appendix B.

more likely than women to have combinations of more life-threatening conditions (e.g., emphysema, respiratory cancer, diabetes, and CVD), whereas women's comorbidities are clustered around conditions that are less likely to be fatal (e.g., conditions that involve pain, asthma, and bronchitis).

Even after we adjust for comorbidities, the presence of health conditions is often associated with larger increases in mortality for men than for women, as can be seen in the last three columns of Table 4 or in the graphs of these coefficients in Figure 6. The top panel of Figure 6 shows all 14 conditions. The bottom panel excludes respiratory and

Figure 6. Marginal Effects of Health Conditions on Two-Year Mortality, Men and Women



digestive cancer, which have large effects on the probability of death. Of the eight conditions that have marginal effects on the probability of death in excess of 0.01, five (stomach cancer, CVD, emphysema, other lung disorders, and respiratory cancer) have larger effects on men’s mortality and three (diabetes, reproductive cancer, and bronchitis) have larger effects on women’s mortality. Although the majority of these differences are not statistically significant (see the last column of Table 4), the results, taken together, suggest that serious health conditions pose a greater threat to men than to women.⁶

6. Consistent with our finding for respiratory cancer, Patel, Bach, and Kris (2004) found that being female is associated with longer survival after surgery and chemotherapy treatments for lung cancer.

The larger threat that chronic conditions pose to men's mortality can be seen more formally by decomposing the difference between female and male mortality rates into prevalence and severity effects, as was done for self-rated health and hospitalizations in the previous two sections. In this sample of 45- to 84 year olds, 2.6% of the women and 4.2% of the men died within two years of the survey, resulting in a difference between women and men of -1.6 percentage points. This 1.6-percentage-point difference can be decomposed into a prevalence effect of -0.2 percentage points, a severity effect of -0.6 percentage points, and a remainder unexplained by health conditions of -0.8 percentage points. Thus, 50% of the sex difference in mortality can be explained by these 14 chronic health conditions, with 25% of it ($-0.02 / -0.08$) explained by men being more likely than women to have the chronic conditions that have larger effects of mortality and 75% of it ($-0.06 / -0.08$) explained by men having a greater probability of dying than women with the same chronic conditions. However, this decomposition provides a misleading estimate of the size of the severity effect. Fully half the severity effect is due to the sex difference in the estimated effect of arthritis on mortality. As is shown in Table 4, arthritis is not associated with male mortality and is *negatively* associated with female mortality, so that it contributes to the severity effect for men. Because arthritis is prevalent, it is heavily weighted in the calculation of the severity effect. When the effects of arthritis on mortality are set to zero for both women and men—which seems sensible, given that we do not think that arthritis really protects women against death—the severity effect drops to -0.2 percentage points, equal to the prevalence effect.

DISCUSSION AND CONCLUSIONS

The results of this study partly explain the paradox with which we started, but do not resolve it entirely. The hypothesis that sex differences in self-rated health can be entirely explained by sex differences in the distribution of conditions is confirmed. The hypothesis that women and men form assessments of their health in different ways is not consistent with either the evidence that women and men with the same health conditions are equally likely to report being in poor health or with our findings that reports of poor health are equally predictive of hospitalization episodes for men and women. However, the evidence for hospitalizations and mortality indicate that men with some specific health conditions are more likely to be hospitalized and to die than are women with the same conditions.

The severity effects that disadvantage men are driven mainly by a small number of smoking-related conditions: CVD (for hospital episodes) and CVD, emphysema, and other lung disorders (for mortality).⁷ An important question is why men with these conditions are more likely than women to be hospitalized and to die. One hypothesis is that because men have had higher rates of smoking throughout their lives than women, they experience more-severe forms of these conditions when they occur.⁸ Although we do not have the data to test this hypothesis, it is consistent with the findings of other research. For example, Pampel (2002) concluded that smoking patterns fully explain the recent narrowing of the

7. Stomach cancer and respiratory cancer, both smoking related, also have larger effects on mortality and hospitalizations for men than for women. However, because these conditions have low prevalence rates, they contribute little to the overall severity effects for mortality or hospitalizations. The link between smoking and stomach cancer is not as well known as the link between of smoking and CVD and respiratory conditions. However, the American Cancer Society (2004) reported that "smoking increases stomach cancer risk, particularly for cancers of the proximal stomach (the upper portion of the stomach closest to the esophagus). The rate of stomach cancer is approximately doubled in smokers."

8. The sex differences in smoking are striking. For example, information on smoking histories from the 1997–2001 NHIS indicated that among all 75 year olds (including those who never smoked), men reported an average of more than 25 years of smoking, in contrast to fewer than 15 years of smoking for women. Years of smoking were lower among those in the younger cohorts, so that among 45 year olds, men had, on average, only 2 more years of smoking over their lives than did women.

mortality differences between men and women, and Valkonen and van Poppel (1997) concluded that 40% of the total sex differences in life expectancy at age 35 could be explained by smoking. If sex differences in the lifetime use of tobacco are generating excess male mortality, then the smaller sex gap in smoking among these younger cohorts may reduce excess male mortality through *both* the prevalence effect and the severity effect.

The remaining question is why, if smoking-related conditions are more likely to result in hospitalizations and death for men, do they not also result in greater reports of poor health for men? A partial answer to this question is that some of these conditions do produce both excess male mortality and higher reports of poor health for men than for women. The results presented in Figure 3 indicate that among those with emphysema and "other lung" problems, men are more likely than women to report poor health. Although sex differences in these coefficients are statistically significant only for lung problems, not for emphysema, they point to the possibility that men experience more-severe forms of these conditions—a possibility that is consistent with the evidence of greater rates of smoking among men presented earlier. However, it is more difficult to explain the finding that men with CVD are equally likely to report poor health but are more likely to experience hospitalizations and to die than are women with CVD.

Another possible answer is that men tend to report health conditions only when they are at more-severe or more-advanced stages of these conditions. For example, a woman with mild angina may report that she has CVD, whereas a man may not do so until he has had a heart attack. This possibility would explain why reported CVD has a larger effect on male than on female mortality. However, for the results for self-rated health to be reconciled with those for mortality, women with milder forms of CVD would have to be as likely to report poor health as men with more-severe forms of CVD. In other words, women would have to be more likely than men to report less-severe forms of conditions *and* to factor milder forms of conditions into their reports of poor health. Although this answer is logically possible, we do not think that it is plausible. If women systematically reported milder forms of all conditions, we would expect to see these conditions having smaller effects on women's mortality. For many of the health conditions we examined, there were no sex differences in the effects on either self-rated health or mortality. This was the case even for conditions, such as diabetes and bronchitis, that can have various degrees of severity.

Additional evidence comes from looking at samples of individuals who died from specific causes (e.g., acute myocardial infarction or stroke) within two years of the survey and examining whether these individuals reported conditions that were associated with those causes of death. If men are more likely than women to underreport conditions, then men who died should have been less likely than women who died to have reported that they had the condition that resulted in their death. The results in Table 5 indicate that this is not the case. Women and men whose causes of death were digestive cancer, diabetes, heart attack, stroke, hypertensive disease, and respiratory cancer were equally likely to have reported the relevant condition in the preceding interview. Men who died of respiratory disease (bronchitis, asthma, emphysema, or chronic obstructive pulmonary disease) were more likely than women who died from respiratory disease to have reported one or more of the conditions leading up to this cause of death. It should be noted that reports of the relevant conditions are often low for both men and women, indicating that chronic conditions may be underreported because of either an unwillingness to disclose health conditions or a lack of knowledge about those conditions (for additional evidence of the underreporting of health conditions in Canada, see Baker, Stabile, and Deri 2004). However, we found no evidence that men report conditions less often than women.

The final explanation is that, for at least some conditions, the symptoms that individuals experience may convey little information about the severity of the disease. This

Table 5. Reports of Chronic Conditions for Those With Specific Causes of Death

Cause of Death	Condition	Women		Men		χ^2 (<i>p</i> value)
		<i>N</i>	% With Condition	<i>N</i>	% With Condition	
Digestive Cancer	Digestive cancer	363	3.31	384	5.21	1.65 (0.199)
Diabetes	Diabetes	177	51.40	141	54.60	0.32 (0.164)
Heart Attack	Cardiovascular disease	544	32.40	669	35.90	1.65 (0.199)
Stroke	Cardiovascular disease	393	33.80	304	35.50	0.21 (0.643)
Hypertensive Disease	Hypertension	122	60.70	93	49.50	2.68 (0.102)
Respiratory Disease	Asthma, bronchitis, or emphysema	267	41.60	319	54.20	9.33 (0.002)
Respiratory Cancer	Respiratory cancer	15	3.91	31	5.51	1.26 (0.261)

Notes: Each row of this table is based on a sample of respondents who died within two years of the survey and had the specific cause of death listed in the first column. The table shows the percentages (of those who died from each cause) who reported the condition listed in the second column. The χ^2 test shows whether the percentages of women and men who reported the condition differ. The causes of death are from the NHIS 72-item ICD-9 recode and are defined as follows: Digestive cancer = malignant neoplasms of digestive organs and peritoneum. Diabetes = diabetes mellitus. Heart attack = acute myocardial infarction. Stroke = intracerebral and other intracranial hemorrhage; cerebral thrombosis and unspecified occlusion of cerebral arteries; cerebral embolism; and "all other and late effects of cerebrovascular diseases." Hypertensive disease = hypertensive heart disease; hypertensive heart and renal disease; and hypertension with or without renal disease. Respiratory disease = bronchitis, chronic and unspecified; emphysema; asthma; and "other chronic obstructive pulmonary diseases and allied conditions." Respiratory cancer = malignant neoplasms of the respiratory and intrathoracic organs.

explanation may be especially relevant for CVD, several components of which can be symptomless. It is possible that women and men who report having CVD experience similar symptoms and therefore report (on average) the same self-rated health when, in fact, the men's disease is more severe (perhaps because of their greater levels of smoking). This explanation is consistent with our finding that men and women with CVD are equally likely to report poor health, that men with CVD are only slightly more likely than women with CVD to experience hospital episodes, and that men with CVD are more than 1.5 times more likely to die than are women who report CVD.

In summary, our results indicate that women and men who have the same health conditions are equally likely to report that they are in poor health. Although the effects of many health conditions on hospitalizations and mortality are similar for men and women, men with some health conditions—notably those that are associated with smoking—are more likely than women with these conditions to experience hospitalizations and death. Overall, the results suggest that the paradox of worse self-rated health for women and higher mortality for men has a fairly straightforward explanation that does not rest on systematic differences in how women and men report their health.

APPENDIX A: DEFINITIONS OF THE VARIABLES

All regressions include a complete set of indicators for age and race (black, white, or "other") and a control for years of education. The chronic health conditions used in the analyses are defined as presented in Table A1 for each of the two survey periods. In coding the conditions, we attempt to make the definitions for conditions from the earlier and later periods as close as possible. For the 1986–1994 surveys, we provide the diagnostic codes that were aggregated to create each condition. The reference periods over which the conditions were assessed are noted for the 1997–2001 period.

Appendix Table A1. Diagnostic Codes That Were Aggregated to Create Each Chronic Health Condition

Condition	Definition
1986–1994 ^a	
Arthritis	Arthritis (101); rheumatism, unspecified (102); gout, including gouty arthritis (103)
Skin cancer	Malignant neoplasms of the skin (119)
Digestive cancer	Malignant neoplasms of the stomach, intestines, colon, or rectum (316)
Reproductive cancer	Malignant neoplasms of the breast (421), female genital organs (422), or prostate (423)
Cardiovascular disease	Rheumatic fever “ever” (501), ischemic heart disease “ever” (502), tachycardia or rapid heart (503), heart murmur (504), other unspecified heart rhythm disorder (505), congenital heart disease “ever” (506), other selected diseases of the heart excluding hypertension (507), cerebrovascular disease “ever” (509), hardening of the arteries “ever” (510), or aneurysm (511)
Hypertension	Hypertension “ever” (508)
Circulation problems (515)	Phlebitis or thrombophlebitis (512), hemorrhoids (514), poor circulation
Bronchitis	Chronic bronchitis (601)
Asthma	Asthma (602)
Emphysema	Emphysema (609)
Other lung disorders	Pleurisy (610), pneumoconiosis and asbestosis (611), tuberculosis (612), or other diseases of the lung (614)
Respiratory cancer	Malignant neoplasms of the lung or bronchus (613) or malignant neoplasms of other respiratory sites (615)
1997–2001	
Headache	Severe headache or facial pain in the past 3 months
Other pain	Pain in the neck, lower back, or joint injury in the past 3 months
Arthritis	Joint pain not due to injury in the past 12 months
Cardiovascular disease	“Ever been told” has had a heart attack, coronary heart disease, angina, another kind of heart condition/heart disease, stroke
Diabetes	“Ever been told” has diabetes
Hypertension	“Ever been told” has hypertension
Circulatory problems	Difficulty with activities because of a “circulatory problem”
Chronic vision problems	Trouble seeing “even when wearing glasses”
Chronic hearing problems	Wears a hearing aid, has “a lot of trouble” hearing, or reports being deaf
Depression	The respondents were asked to report whether they felt sad, nervous, restless, hopeless, that everything was an effort, or worthless over the past 30 days. Individuals who reported having at least one of these feelings either “all of the time” or “most of the time” were coded as depressed.
Bronchitis	Has been told has chronic bronchitis in the “past 12 months”
Emphysema	“Ever been told” has emphysema
Asthma	“Ever been told” has asthma

(continued)

(Appendix Table A1, continued)

Condition	Definition
1997–2001 (cont.)	
Other lung problems	Difficulty with activities because of a “lung or breathing problem”
Skin cancer	“Ever diagnosed” with skin cancer
Stomach cancer	“Ever diagnosed” with stomach cancer
Reproductive cancer	“Ever diagnosed” with reproductive cancer
Respiratory cancer	“Ever diagnosed” with respiratory cancer

^a12-month reference period unless noted otherwise.

APPENDIX B: BIAS CORRECTIONS FOR MORTALITY ESTIMATES

The equation to be estimated is

$$D = C_0\beta_0 + C_1\beta_1 + \dots + C_M\beta_M + \epsilon = C\beta + \epsilon, \tag{A1}$$

where it is assumed that ϵ , is *i.i.d.* with $E[\epsilon | C] = 0$. C_0 is a matrix of demographic variables (age, education, and race) that are observed for all individuals. For $j = 1 \dots M$, C_j is an $N \times k_j$ matrix of 0/1 variables that indicate whether the individual has each of the conditions included in condition list j , where there are a total of M condition lists. Assume that the number of individuals who are asked about the conditions in each list is equal to $n = N / M$ and that the total number of explanatory variables is $\sum_{j=0}^M k_j = k$. The matrix C is censored, so that for individuals who are assigned to condition list j , only variables in C_0 and C_j are observed. The outcome D is observed for all individuals.

Define a matrix Z that is equal to C but with censored elements set to 0, so that

$$C = \begin{bmatrix} C_{01} & C_{11} & \dots & C_{M1} \\ C_{02} & C_{12} & \dots & C_{M2} \\ \cdot & \cdot & \dots & \cdot \\ C_{0M} & C_{1M} & \dots & C_{MM} \end{bmatrix} \text{ and } Z = \begin{bmatrix} C_{01} & C_{11} & 0 & \dots & 0 \\ C_{02} & 0 & C_{22} & \dots & 0 \\ \cdot & \cdot & \dots & \cdot & \cdot \\ C_{0M} & 0 & 0 & \dots & C_{MM} \end{bmatrix},$$

where C_{ij} is an $n \times k_j$ matrix of information for individuals who are assigned to condition list j .

In what follows, it will be useful to have probability limits of the matrices $(C'C) / N$, $(Z'C) / n$, and $(Z'Z) / n$. Assume that

$$plim_{N \rightarrow \infty} \frac{C'C}{N} = \Sigma = \begin{bmatrix} \Sigma_{00} & \Sigma_{01} & \dots & \Sigma_{0M} \\ \Sigma_{10} & \Sigma_{11} & \dots & \Sigma_{1M} \\ \cdot & \cdot & \dots & \cdot \\ \Sigma_{M0} & \Sigma_{M1} & \dots & \Sigma_{MM} \end{bmatrix}. \tag{A2}$$

Then

$$plim_{n \rightarrow \infty} \frac{\mathbf{Z}'\mathbf{C}}{\mathbf{n}} = \Sigma_{\mathbf{ZC}} = \begin{bmatrix} M \Sigma_{00} & M \Sigma_{01} & \dots & M \Sigma_{0M} \\ \Sigma_{10} & \Sigma_{11} & \dots & \Sigma_{1M} \\ \cdot & \cdot & \dots & \cdot \\ \Sigma_{M0} & \Sigma_{M1} & \dots & \Sigma_{MM} \end{bmatrix} \tag{A3}$$

and

$$plim_{n \rightarrow \infty} \frac{\mathbf{Z}'\mathbf{Z}}{\mathbf{n}} = \Sigma_{\mathbf{ZZ}} = \begin{bmatrix} M \Sigma_{00} & \Sigma_{01} & \Sigma_{02} & \dots & \Sigma_{0M} \\ \Sigma_{10} & \Sigma_{11} & \mathbf{0} & \dots & 0 \\ \Sigma_{20} & \mathbf{0} & \Sigma_{22} & \dots & 0 \\ \cdot & \cdot & \cdot & \dots & 0 \\ \Sigma_{M0} & 0 & 0 & \dots & \Sigma_{MM} \end{bmatrix}. \tag{A4}$$

Finally, assume that there is a supplemental data set that contains N_s observations on all variables, with the $N_s \times k$ data matrix denoted \mathbf{C}_s . Assume that

$$plim_{N_s \rightarrow \infty} \frac{\mathbf{C}'_s \mathbf{C}_s}{\mathbf{N}} = \Sigma.$$

The starting point of the bias correction is the OLS estimator of β using \mathbf{Z} in place of \mathbf{C} , that is, $\tilde{\beta} = (\mathbf{Z}'\mathbf{Z})^{-1} \mathbf{Z}'\mathbf{D}$. The bias-corrected estimate of β shown in Eq. (6) is

$$\hat{\beta} = \hat{\Sigma}_{\mathbf{ZC}}^{-1} \hat{\Sigma}_{\mathbf{ZZ}} \tilde{\beta}. \tag{A5}$$

The estimate of $\hat{\Sigma}_{\mathbf{ZZ}}$ is obtained from the censored data as $(\mathbf{Z}'\mathbf{Z}) / n$. As in Eq. (A4), the lower right portion of this matrix is block diagonal, with estimates of the matrices Σ_{jj} for $j = 1 \dots M$ along the diagonal. The elements of these matrices are equal to

$$\hat{\Sigma}_{jj}(i, k) = P_{ik}^j,$$

where P_{ik}^j is the fraction of individuals who are assigned to condition list j who have both the i th and k th condition within that list. When i is equal to k , this is simply equal to the prevalence of the condition.

The matrix $\Sigma_{\mathbf{ZC}}$ is estimated using the censored and supplemental samples. Specifically, cross-products that are based on within-condition-list terms are taken from the censored sample and are identical to the corresponding blocks in the estimate of $\Sigma_{\mathbf{ZZ}}$. The off-block-diagonal terms that represent cross-products that cross condition lists are estimated as follows:

$$\hat{\Sigma}_{jl}(i, k) = P_{ik} = P_i P_k + \rho_{ik} \sqrt{P_i P_k (1 - P_i)(1 - P_k)}, \tag{A6}$$

where P_i and P_k equal the fraction of the censored sample that has condition i and k , respectively, and ρ_{ik} is the correlation between condition i and k that is estimated from the supplemental sample. An alternative method of obtaining $\hat{\Sigma}_{jl}(i, k)$ would have been to compute the fraction of the supplemental sample that has both condition i and k . However, for some pairs of rare conditions, the fraction of the supplemental sample that had both condition i and condition k exceeded the fraction of the censored sample that

had either i or k . The use of Eq. (A6), which draws only the correlation coefficient between i and j from the supplemental sample, prevents this from occurring.

The proof that $\hat{\beta}$ is consistent is straightforward. Substituting the formula for $\hat{\beta}$ into Eq. (A5), we get

$$\hat{\beta} = \hat{\Sigma}_{ZC}^{-1} \frac{Z'C}{n} \beta + \hat{\Sigma}_{ZC}^{-1} \frac{Z'\epsilon}{n},$$

the first term of which has a probability limit of β and the second term of which has a probability limit of 0.

The estimate of the variance-covariance matrix for $\hat{\beta}$ is

$$\frac{\sigma^2}{n} \hat{\Sigma}_{ZC}^{-1} \hat{\Sigma}_{ZZ} \hat{\Sigma}_{ZC}^{-1}, \quad (A7)$$

where σ^2 is the variance of ϵ . A consistent estimate of σ^2 , denoted $\hat{\sigma}^2$, is obtained by starting with the estimate of the error variance using the error-ridden estimate $\hat{\beta}$. Specifically, let

$$\epsilon = D - Z\hat{\beta} = [I - Z(Z'Z)^{-1}Z']D = M_Z[C\beta + \epsilon].$$

Then

$$\frac{\tilde{\epsilon}'\tilde{\epsilon}}{N-k} = \tilde{\sigma}^2 = \frac{\beta'C'M_ZC\beta}{N-k} + \frac{2\epsilon'M_ZC\beta}{N-k} + \frac{\epsilon'M_Z\epsilon}{N-k}$$

and

$$E[\tilde{\sigma}^2] = \frac{\beta'C'M_ZC\beta}{N-k} + \sigma^2. \quad (A8)$$

The estimate $\hat{\sigma}^2$ is obtained by subtracting the first term on the right-hand side of Eq. (A8) from $\tilde{\sigma}^2$. Substituting in the definition of $M_Z = I - Z(Z'Z)^{-1}Z'$ yields

$$\hat{\sigma}^2 = \tilde{\sigma}^2 - \frac{N}{N-k} \left[\hat{\beta}' \hat{\Sigma} \hat{\beta} - \frac{1}{M} \hat{\beta}' \hat{\Sigma}_{ZC} \hat{\Sigma}_{ZZ}^{-1} \hat{\Sigma}_{ZC} \hat{\beta} \right],$$

where $\hat{\Sigma}$ is constructed from $\hat{\Sigma}_{ZC}$ —specifically, where $\hat{\Sigma}$ equals $\hat{\Sigma}_{ZC}$ with all elements of the first k_0 rows divided by M (compare Eqs. (A2) and (A3)).

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