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Mortality and Lifetime Income: Evidence from U.S. Social Security Records

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Studies of the empirical relationship between income and mortality often rely on data aggregated by geographic areas and broad population groups and do not distinguish between disabled and nondisabled persons. This paper investigates the relationship between individual mortality and lifetime income with a large microdatabase of current and former retired participants in the U.S. Social Security system. Logit models by gender and race confirm a negative relationship. Differences in age of death between low and high levels of lifetime income are on the order of two to three years. Income-related mortality differences between blacks and whites are largest at low-income levels, but gender differences appear to be large and persistent across income levels. [JEL C40, C67, D31, H20, I38]
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Among the most critical issues in designing a public pension scheme are the redistributive aspects of the system. Pension plans, particularly defined benefit plans, can redistribute wealth across population groups in both intended and unintended ways. Some types of redistribution are

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unavoidable, such as the higher benefits paid to an individual who experiences an unusually long lifetime relative to one who dies earlier due to illness or injury. This kind of inherent probabilistic transfer can be offset through disability and survivor provisions and through conditions that attempt to ensure that *expected* returns are more nearly equal. By contrast, redistribution across income groups is often included intentionally into the design of the system, by making replacement rates a function of income, establishing minimum or maximum benefits, or other progressivity provisions (Congressional Budget Office, 2006).

Evaluations of the degree of progressivity of a pension system require specific knowledge of the relationship between life expectancy and income. If wealthy individuals tend to live longer, they will experience higher rates of return and greater present values of net benefits than would be implied by system parameters alone. Although abstraction from some aspects of differential mortality risk may be purposeful (such as by not using gender- or race-specific mortality tables), abstraction from the differential risk between low- and high-income groups would seem perverse. This paper examines the relationships of life expectancy to lifetime income, providing new information to help future policymakers and analysts.

The calculation of initial benefits in the U.S. Social Security system, for example, makes no attempt to reflect systematic differences across the population in mortality risk. Individuals of the same age with identical earnings histories, and therefore with identical contribution amounts, receive the same initial benefit at retirement irrespective of differences in expected length of life. One consequence is that unintended redistribution can occur that depends on differences in life expectancy related to such factors as income, gender, and race. Milton Friedman pointed this out over 30 years ago, noting that higher-income people have higher life expectancy and therefore receive benefits for a longer period of time (Friedman, 1972). Proposals to reform the U.S. Social Security system, including proposals for annuitization of individual account provisions, generally ignore mortality risk and could cause significant transfers from groups with higher mortality risk to those with lower risk.

Economists, sociologists, demographers, and health professionals have all studied the relationship between life expectancy and socio-economic status, and the literature has grown dramatically since the oft-cited study of Kitagawa and Hauser (1973). Nevertheless, the literature in this area is relatively young and appropriate data are scarce. Most early research and some recent work rely on data aggregated by geographic areas and broad population groups, often categorizing people into income classes based on a single year of income. Such an approach cannot distinguish between populations that are temporarily and those that are permanently in a certain income class, obscuring the distinct effects of permanent and transitory income. Moreover, most available databases do not distinguish disabled and nondisabled persons. Because disabled persons have far higher mortality than the general population (Zayatz, 1999) and because their lifetime income is likely to be affected by their health status, combining the two populations may result in seriously misleading implications of the income-mortality relationship for social security calculations.

This paper analyzes empirically the seldom-investigated relationship between mortality and individual-specific lifetime income with data that largely avoid the aforementioned problems. The data are administrative records from the Social Security Administration's (SSA) 2002 One Percent Continuous Work History Sample (CWHS) that has actual covered earnings information spanning the period 1937 to 2002. The file contains month and year of death information for deceased persons covered by the social security program, along with race and gender indicators. Our data set also includes all of the benefit records from SSA's Master Beneficiary Record (MBR) file that are associated with the CWHS records. The benefit records contain additional death information and allow us to separate the mortality experience of retired from disabled workers.

Our results give strong empirical support to a negative relationship between individual lifetime income and mortality for persons who have survived until age 62. For black and white males and females the difference in age of death between low and high lifetime income is on the order of two to three years. Workers with positively trended earnings over their work life may live an additional six to 18 months relative to those with declining earnings. Income-related mortality differences between blacks and whites are largest at low-income levels, particularly for males, and narrow substantially at higher income levels. On the other hand, gender differences in mortality appear to be large and persistent across income levels.

I. Previous Research on Income and Mortality

Prior research has generally reported a negative relationship between income and mortality. That finding is not universal, however, particularly in the case of older males and females, the focus of this paper. The seminal work by Kitagawa and Hauser (1973), for example, found a strong negative effect of current income on mortality for persons under 65 years of age but a much smaller and sometimes positive effect for persons over 65, a finding confirmed by Rosen and Taubman (1979) with later data. Hadley and Osei (1982); House, Kessler, and Herzog (1990); and Sorlie, Backlund, and Keller (1995) found a positive or weak relationship between current income and mortality at older ages. A difficulty with earlier results is that current income is a poor measure of lifetime resources, a more appropriate concept for differentiating mortality probabilities for that population group. This has led some researchers to prefer proxies like education (Brown, 2002) or even components of current income (Krueger and others, 2003). Current income also raises concerns about the direction of causality, particularly when the data do not allow control for health status.

By contrast, in this paper we use a direct measure of lifetime resources—lifetime earnings—for each individual in the sample and exclude individuals with a serious disability condition. Lifetime earnings, unlike current income, captures the value of resources earned over prime working years and, for the age groups examined here, is not subject to concerns over the direction of causality. As indicated above, we estimate a strong negative effect on the probability of death for older men and women.

Obviously, other factors also affect mortality. Mortality may vary by occupation, although most of the occupational differences in mortality are explained by income or education (Wilmoth and Dennis, 2006). Several researchers have assessed the comparative roles of race and other socioeconomic variables in determining death hazards, but agreement is lacking. Some authors find that black-white differences in mortality are fully explained by socioeconomic status but others find that socioeconomic status reduces but does not eliminate black-white mortality differences (Wilmoth and Dennis, 2006). This paper provides new evidence on the income-race-mortality nexus.

II. Data and Variable Definitions

The primary data sources used here are the SSA's 2002 CWHS and MBR. The "active" portion of the CWHS (that is, individuals with some covered earnings) is a historical record of social security earnings for over three million current or former workers in covered employment spanning the period 1937 to 2002. Each record has date of birth and race and sex indicators. When an individual files a claim for benefits based on her or his previous earnings history, SSA opens a claim account in the MBR file, the principal official record of historical benefit payments. If a worker attains beneficiary status and subsequently dies, the date of death is recorded in the MBR. Death information in the MBR is considered to be of very high quality (Aziz and Buckler, 1992).

The administrative records have several advantages over data sets used in previous analyses of the determinants of mortality. The coverage is unparalleled, with the number of records many times larger than in any of the surveys with information on date of death. The CWHS has individual longitudinal information on earned income, as well as demographic information (age, race, sex), that is matched to mortality data. These data obviate aggregation or one-time measurement of the income variable that is of primary interest.[1]

The CWHS also has limitations for estimating mortality models. Most important for our purpose is that it contains information only on income earned in covered employment. Also, family background information that may impact on mortality rates is not included. Finally, race information in

[1] Hoyert, Singh, and Rosenberg (1995) summarize the principal data sets available to evaluate mortality and socioeconomic status.

the administrative data permits only a three-way classification (blacks, whites, others).[2]

For empirical analysis we selected all retired-worker beneficiary records with birth years 1900 to 1942, who claimed benefits at age 62 or later. The birth year range allows us to observe a complete life for early cohorts and a nearly complete work life for later cohorts. Because our benefit/death information extends through 2004, the oldest observed age in our sample is 104. Accounting for edits and deletions for anomalous cases, our sample has nearly 550,000 observations.[3] Some characteristics of our sample are shown in Table 1.

Nearly half of retired-worker beneficiaries in our sample have died. The percentages are higher for males than females and higher for blacks than whites. For retired-worker beneficiaries who have already died, the median female lived about four years longer than the median male and the median white lived about one to two years longer than the median black.

Lifetime earnings are measured for each individual as the sum of real earnings (2005 dollars) over ages 35 to 60 (ages 37 to 60 and 36 to 60 for birth years 1900 and 1901, respectively).[4] As seen in Table 1 males earned more than females and whites earned more than blacks. The relationship between age of death and lifetime earnings for deceased male retired workers in our sample is illustrated in Figure 1 (a similar but less steeply sloped pattern arises for females). The figure shows three five-year birth cohorts, selected to represent patterns across the birth cohorts in our sample, with each five-year birth cohort sorted into deciles of lifetime earnings. The figure plots the

[2]Since 1980, race information collected in the administrative records has expanded to include five race/ethnic responses: white (not Hispanic), black (not Hispanic), Hispanic, Asian or Pacific Islander, and American Indian or Alaskan Native. Prior to 1980, only the three-way classification was allowed. The vast majority of records analyzed in this paper would not, of course, benefit from the expanded classification.

[3]One anomaly in our sample occurs with records that show no recent work activity, no benefits, insufficient quarters of coverage for benefit eligibility, and no death information. These records accounted for less than 9 percent of the sample before edits. The vast majority of these records are missing death information in our file due to state agreements with SSA on the use of death reports. We deleted these records on the grounds that their exclusion would be less distorting than their inclusion. The exclusion has no effect on our later econometric estimates that focus on beneficiaries.

[4]Earned income in the CWHS consists of lump-sum taxable earnings for 1937 to 1950, annual taxable earnings from 1951 to 1977, and total (uncapped) earnings from 1978 to 2002. We use these earnings data to develop a summary measure of real lifetime earnings following three adjustments to the data: (1) we prorated the 1937 to 1950 lump-sum under the assumption that, between the reported first and last years of employment prior to 1951, earnings grew at 1 percent per year of age beyond the economy-wide growth in wages for males and at 0.5 percent for females, using the average annual earnings between 1937 and 1950 from Myers (1993); (2) for individuals who earned the taxable maximum during years 1937–80, we imputed earnings above the maximum using a Tobit model of earnings under the assumption that earnings are lognormally distributed; (3) earnings (actual or imputed) in each year were indexed to 2005 using the consumer price index (CPI).

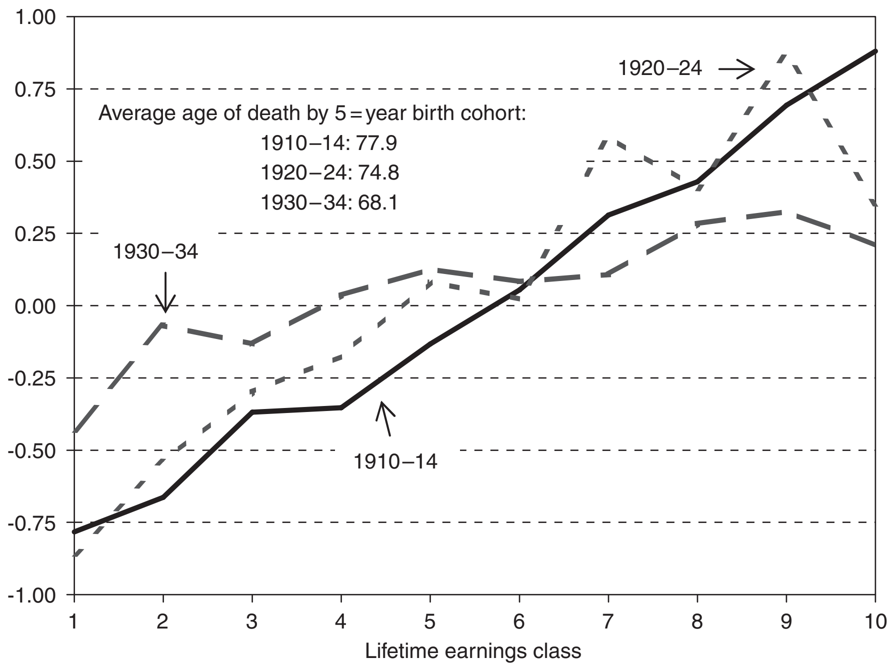
Table 1. Characteristics of Retired Worker Beneficiaries

		Females			Males		
Category	Sample	White	Black	Other	White	Black	Other
Sample (%)	100.0	38.5	3.7	1.1	50.6	4.4	1.7
Dead (%)	49.0	43.6	45.4	24.4	54.2	59.0	37.5
Median age of death	77.0	80.0	78.0	79.0	76.0	75.0	77.0
Median lifetime earnings	517,158	269,574	236,917	233,227	921,999	476,051	413,088

Source: The data are from a merge of matched records from the Continuous Work History Sample (CWHS) and the Master Beneficiary Records (MBR) of the Social Security Administration.

Note: This table presents the characteristics for a sample of retired worker beneficiaries aged 62 and over. Lifetime earnings are defined as aggregate labor earnings between ages 35 and 60, measured in 2005 dollars. Sample size = 548,681.

Figure 1. Deviations from Average Age of Death for Male Retired Workers



Source: Authors' calculations using matched records from the CWHS and MBR of the SSA.

Note: This figure displays deviations from the average age of death for male retired workers for three different five-year birth cohorts classified by within-cohort deciles of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars).

differences between the average age of death in each earnings class and the average age of death across the earnings classes, thereby adjusting for scale differences across birth cohorts. For the 1910–14 birth cohorts, difference in age of death between highest and lowest earnings deciles is about 20 months; for the 1920–24 cohorts, the difference is about 15 months and about 8 months for the most recent cohorts. The figure clearly depicts a positive relationship between lifetime earnings and length of life. Later we will examine in some detail the importance of this measure of "permanent income" for mortality outcomes.

We are also interested in the relationship between the *pattern* of earnings over the work cycle and life expectancy. Many workers exhibit low adherence to the workforce due to limited skills or poor health but others, for similar reasons, may have steady workforce attachment but achieve little progress in real earnings. On the other hand, workers who continue to improve their earnings positions are likely to be more skilled, in better health, and generally more productive. The latter group may also have longer life expectancies. We attempt to capture these effects by classifying workers according to whether the trend in their real earnings over prime working ages has risen, remained flat, or fallen.[5] Specifically, consider an individual worker's average real earnings during the three equal age periods 37 to 44, 45 to 52, and 53 to 60 and label the periods AE1, AE2, and AE3, respectively. The trend in a worker's real earnings can be measured as

$$Trend = (AE3 - AE1)/(AE1 + AE3).$$

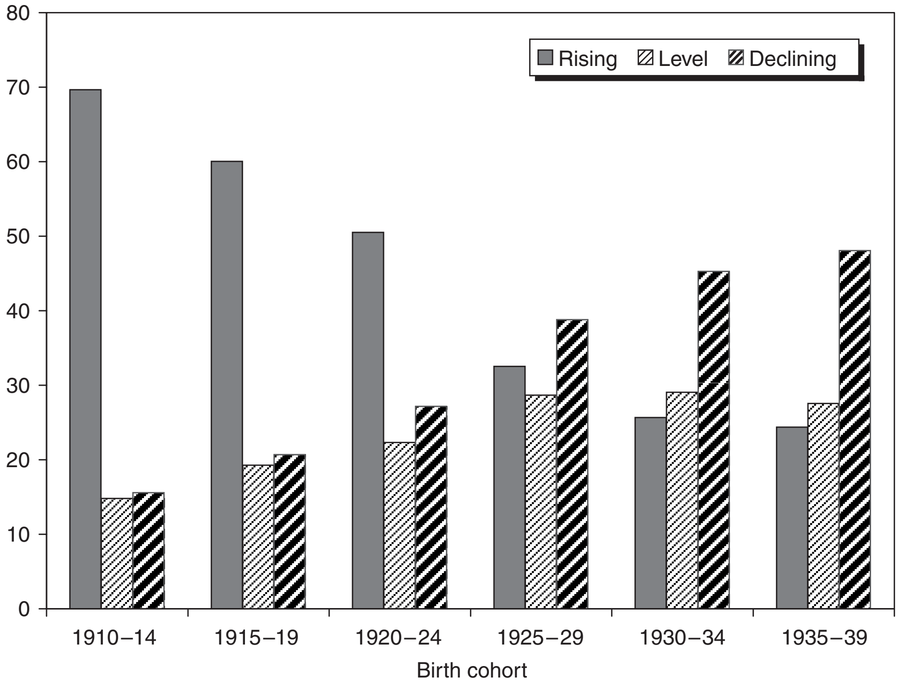
We classify workers according to earnings trend as follows:

Declining if trend $< -1/9$,
Flat if $-1/9 \leq$ trend $\leq 1/9$,
Rising if trend $> 1/9$.

Figure 2 shows a steady decline in the percentage of male retired workers with a rising trend and a steady rise in the percentage with a declining trend. A similar but more modest change in trend has occurred for females. Figure 3 illustrates the relationship between cumulative age of death and the trend in earnings for deceased male retired workers. The age-of-death distribution for male workers with a declining earnings trend lies above the distribution for workers with rising or flat trends. The difference in median age of death between those with rising trends and those with declining trends is about four years.

[5]The earnings classification scheme described below is based on work performed on the 1972 CWHS and reported in the Committee on Finance of the U.S. Senate and the Committee on Ways and Means of the U.S. House of Representatives (1976, pp. 75–81).

Figure 2. Distribution of Male Retirees by Trend in Earnings
(In percent)



Source: Authors' calculations using matched records from the CWHS and MBR of the SSA.
Note: This figure displays the shares of workers within five-year birth cohorts with rising, level and declining labor earnings.

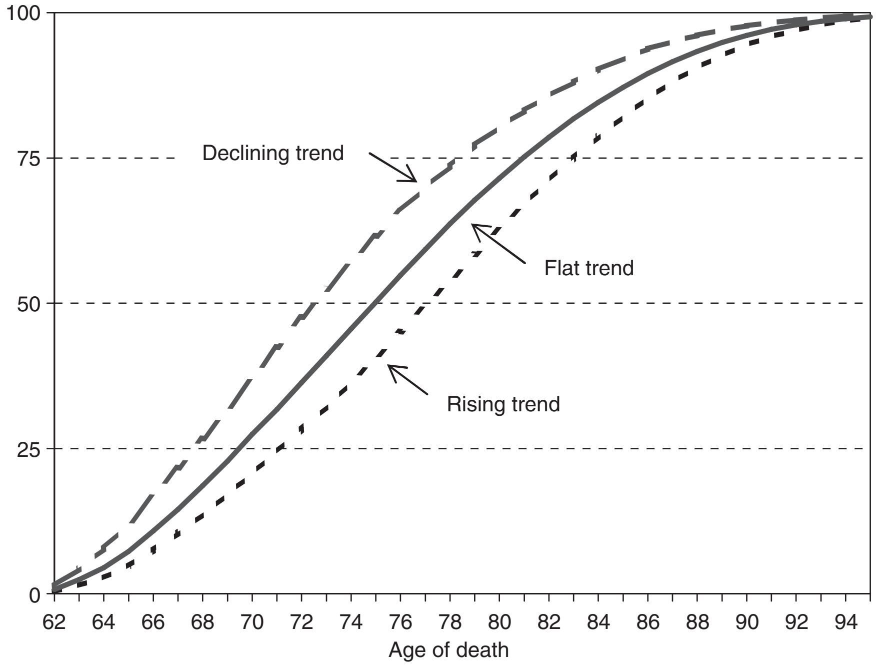
III. Empirical Mortality Models

This section estimates mortality models for white and black male and female retired workers. The unit of observation is an individual year of a worker's life and the dependent variable is the binary realization of the ‘‘hazard rate,'' indicating whether or not death took place in that year, given that the individual was alive at the end of the previous year. We estimate the hazard directly using the logit probability model so that variables with positive coefficients are associated with higher death rates and shorter expected lifetimes.

Logit models have been used extensively in survival analysis generally and have good characteristics for this task.[6] Explanatory variables in our models include (log of) real lifetime earnings, age, year of birth, and indicators of rising and flat earnings trends (declining trend is omitted). (See Appendix I for a discussion of the empirical effect of including the income variables in the mortality model.) In contrast to cross-sectional or aggregate

[6]See, for example, Efron (1988) or Allison (1995, Chapter 7).

Figure 3. Cumulative Distribution of Age of Death of Male Retirees by Trend in Earnings
(In percent)



Source: Authors' calculations using matched records from the CWHS and MBR of the SSA.
Note: This figure presents the cumulative distribution of age of death for male retirees classified by whether their labor earnings exhibited a declining, flat, or rising trend.

data sets, our social security database enables us to separately measure the impact of cohort and age. The cohort variable measures the trend in longevity, but the age variable measures variation in mortality across ages within a cohort. In this specification, real earnings will vary both across and within cohorts.[7] The birth year variable is designed to capture trends in mortality unrelated to the growth in real income over time. The earnings variable is specified as a cubic polynomial in order to allow flexibility in the way earnings affect mortality.

Retired Male Workers

Table 2 displays our parameter estimates for male retired workers (summary statistics for all the logit models are provided in Appendix II). Most

[7]As noted above, the CPI is used to deflate nominal earnings. Duggan and Gillingham (1999) examine potential errors in that index series and the potential impact of those errors on social security finances.

Table 2. Estimated Mortality Hazard Models: Male Retired Workers 62 and Over

Variable	White Males	Black Males
Intercept	−9.991	−9.400
	(0.046)	(0.142)
Age	0.095	0.088
	(0.000)	(0.001)
Birth year	−0.017	−0.006
	(0.000)	(0.001)
Lifetime earnings (LE)		
LE	−0.341	−0.406
	(0.023)	(0.064)
LE^2	0.067	0.079
	(0.004)	(0.010)
LE^3	−0.003	−0.004
	(0.000)	(0.000)
Trend in earnings		
Rising	−0.214	−0.195
	(0.007)	(0.022)
Flat	−0.101	−0.129
	(0.009)	(0.029)

Note: This table reports the estimated parameters of logit mortality hazard models for white and black male retired workers, where the probability of death is a function of age, birth year, lifetime earnings (LE), defined as aggregate labor earnings between ages 35 and 60, measured in 2005 dollars, and whether earnings were declining, rising, or flat during the worker's career. Estimated standard errors are in parentheses.

coefficients are highly significant, reflecting in part the large sample sizes.[8] The probability of death rises with age and falls with birth year. The logit parameters measure the marginal impact of independent variables on the log odds of death; that is, on $\ln[p/(1-p)]$ where p is the hazard rate—the probability of death. Thus, the coefficient of 0.095 in the first column of Table 2 indicates that an additional year of age increases the odds of death by 9 to 10 percent for retired white male workers. The impact of age on p is given by $0.095p(1-p)$. For a white male born in 1920 with the median lifetime earnings from Table 1 and a rising income trend, the estimated probability of death rises from about 1.4 percent at age 65 to 8.7 percent at age 85, but the marginal effect of an additional year of life rises from about 0.1 percent to about 0.7 percent.

The probability of death falls for more recent birth cohorts, presumably reflecting trends in health care and nutrition. For white and black males born

[8]The basic sample consists of 277,571 white males and 24,160 black males. In person-year format, there are over 3.8 million white-male observations and over 317,000 black-male observations.

in 1930, for example, the odds of death are 16 and 5 percent lower, respectively, than for retirees born in 1920, ceteris paribus. The trend in earnings also has a negative effect on mortality. Relative to a declining trend, a rising trend reduces the odds of mortality by about 20 percent and a flat trend reduces the odds by 10 to 13 percent. In each case, mortality is significantly negatively related to the log of lifetime earnings, a result that we explore further below.

Retired Female Workers

For many female workers, the primary source of earned income may not be their own earnings but that of their spouses. In these cases, spousal earnings, in addition to a worker's own earnings, may better capture the effects of lifetime income on mortality. Unfortunately, social security records do not identify marital status per se so that pairing all married workers is not possible. Yet, the records do identify beneficiary couples when an individual receives a benefit based on their own earnings history and an auxiliary benefit based on the earnings of their current or former spouse. These dual beneficiary cases are common among female retirees—about half of female retired workers are dual beneficiaries.[9] For these cases, we are able to match the actual earnings history of the (current or former) spouse with the beneficiary's record and thereby include a second known source of income in the female mortality model.[10]

Female dual beneficiaries are current or former retired-worker beneficiaries who also received either a spousal or widow's benefit based on the earnings of a current or former spouse. Female nonduals are current or former retired-worker beneficiaries with no auxiliary benefit.[11] Duals and nonduals differ considerably in their income and work experience, as shown in Table 3. The median earner nondual white female had nearly three times the earnings of the median dual white female; for blacks, the ratio was nearly two and a half to one. Duals had seven to nine years fewer years of market work than nonduals due at least in part to child-rearing. Finally, the median-earner spouse of white female duals earned nearly seven times that of the dual whereas the median-earner spouse of black females had nearly four times the

[9]Some unknown proportion of female retired worker beneficiaries may be married but not receive an auxiliary benefit because their own benefit exceeds the auxiliary benefit amount. For those individuals, benefit records do not reveal marital status. Male retirees may also be dual beneficiaries but the number is very small.

[10]The lifetimes income of spouses were computed in the same manner as for the primary worker, described in section III. That is, we first impute above-cap earnings for those earning the maximum and then sum their real earnings over ages 35 to 60.

[11]For the empirical analysis, we have 123,979 nonduals and 94,945 dual beneficiaries (about 95 percent of the white and black females represented in Table 1). We omitted about 10 percent of duals who were mainly divorced wives and widows, who were not the primary claimants on a dual account.

Table 3. Income and Work Years of Dual and Nondual Female Beneficiaries

	Duals ($N = 94{,}945$)				Nonduals ($N = 123{,}979$)	
	White		Black		White	Black
	Own	Spouse	Own	Spouse		
Lifetime earnings	150,749	1,013,750	130,029	501,740	420,491	310,227
Years with positive income	20		22		29	29

Source: Authors' calculations using matched records from the Continuous Work History Sample (CWHS) and the Master Beneficiary Records (MBR) of the Social Security Administration.

Note: This table presents median lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars) and years with positive labor earnings between ages 35 and 60 for female beneficiaries classified by race and whether or not they are entitled to benefits under both their own and their spouses work history (duals and nonduals).

earnings of the dual (about 5 percent of black female retired workers are duals).

Table 4 presents parameter estimates for mortality models for the nondual and dual retired female workers. Own lifetime earnings has a significantly negative effect on mortality for both dual and nondual females. For white female duals, lifetime earnings of their spouses also has a significantly negative effect on mortality; the effect is insignificant for black female duals. The variables measuring the trend in earnings have a significantly negative effect on mortality for all four groups represented in Table 4. In the dual equations, birth year has a positive coefficient, in contrast to the birth-year effect in the other mortality equations. This trend could reflect intertemporal changes in the relationship between dual-beneficiary status and health status.[12]

IV. How Large Are Income-Related Differences in Life Expectancy?

In this section we use the estimated mortality models to compare survivor rates and life expectancies at different income levels for retired workers. As backdrop for this analysis, Figure 4 has the cumulative distributions of lifetime income for black and white males and females. The black and white female distributions are quite similar whereas the distribution for black males lies well below that for white males.[13] Almost 80 percent of black males fall below the median value for white males and only 26 percent of white males

[12]This could be investigated using a latent-variable sample-selection model but we chose not to as it is not central to the objective of this paper.

[13]For females, the distributions of dual earnings for blacks and whites are quite similar while the distribution of nondual earnings for whites lies somewhat to the right of that for blacks. Figure 4 combines duals and nonduals.

Table 4. Estimated Mortality Hazard Models: Female Retired Workers

	White		Black	
Variable	Duals	Nonduals	Duals	Nonduals
Intercept	−12.883	−10.324	−12.369	−9.311
	(0.087)	(0.056)	(0.396)	(0.157)
Age	0.129	0.100	0.119	0.085
	(0.001)	(0.001)	(0.003)	(0.001)
Birth year	0.003	−0.015	0.014	−0.012
	(0.001)	(0.001)	(0.003)	(0.002)
Lifetime earnings (LE)				
LE	−0.180	−0.205	−0.451	−0.179
	(0.028)	(0.030)	(0.144)	(0.072)
LE^2	0.039	0.046	0.086	0.048
	(0.005)	(0.005)	(0.024)	(0.012)
LE^3	−0.002	−0.002	−0.004	−0.002
	(0.000)	(0.000)	(0.001)	(0.001)
Trend in earnings				
Rising	−0.263	−0.273	−0.188	−0.289
	(0.010)	(0.011)	(0.042)	(0.030)
Flat	−0.125	−0.116	−0.163	−0.170
	(0.017)	(0.015)	(0.076)	(0.046)
Lifetime earnings of current/former spouse (LESP)				
LESP	−0.342		−0.086	
	(0.049)		(0.155)	
$LESP^2$	0.057		0.024	
	(0.007)		(0.024)	
$LESP^3$	−0.002		−0.001	
	(0.000)		(0.001)	

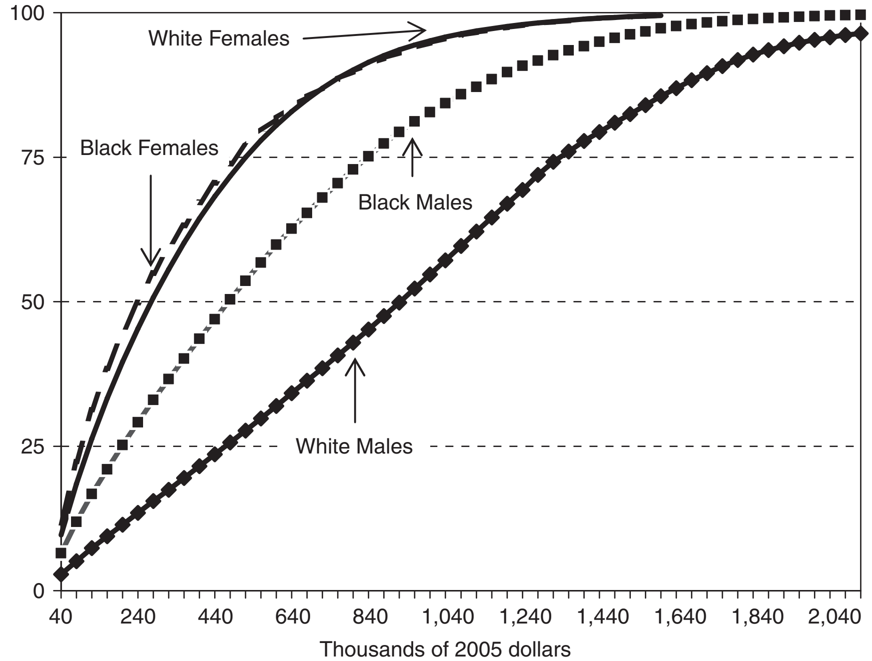
Note: This table reports the estimated parameters of logit mortality hazard models for white and black retired female workers, estimated separately for females who are and are not entitled to a benefit based on their spouse's work history. The probability of death is a function of age, birth year, own lifetime earnings (LE) and, in the case of duals, spouses' lifetime earnings (LESP), each measured as aggregate earnings over ages 35 to 60 expressed in 2005 dollars, and whether their own or, where applicable, their spouse's earnings were declining, flat or rising during the worker's career. Estimated standard errors are in parentheses.

fall below the median value for black males. The shapes of the male distributions are also important when comparing changes in earnings between blacks and whites.

Black and White Males

Figure 5 shows survivor rates, the predicted proportions of age-62 workers still alive at subsequent ages, for hypothetical white male retired workers at the 10th, 50th, and 90th percentiles of the white male earnings distribution,

Figure 4. Cumulative Distribution of Lifetime Earnings for Retired Workers Born 1900–42
(In percent)



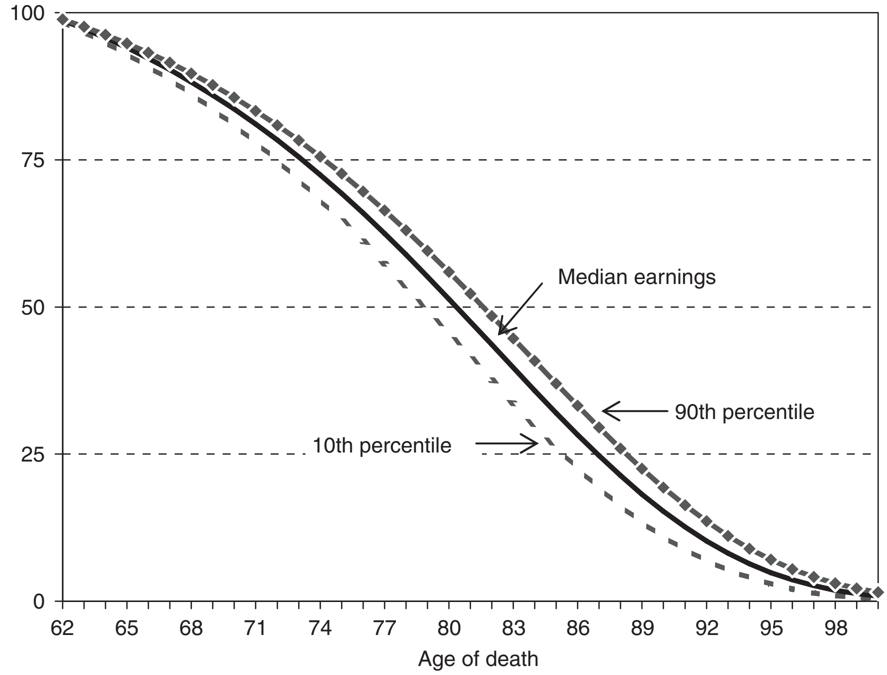
Source: Authors' calculations using matched records from the CWHS and MBR of the SSA.
Note: This figure displays the cumulative distribution of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars) for retirees classified by sex and race.

computed with the coefficient estimates in Table 2. The hypothetical workers were born in 1920 and have a flat real earnings profile. The predicted median age of death for the white male median earner is 80.3 years. For the low and higher earners, the predicted median ages of death are 78.8 and 81.6, respectively (Table 5). For hypothetical workers with a positive earnings profile the median ages of death are about a year higher at each earnings level.

Figure 6 has survivor rates for similarly defined hypothetical black male retired workers. The median age of death for the black median-earner retired worker is 77.8, two and a half years lower than that of the white median earner. For low and high earners, the median ages of death are 76.8 and 80.3, respectively. The age-of-death difference between the lowest and highest earners is over three years (42 months). For black workers a positive trend results in about a half year greater life expectancy.

Differences in the black-white earnings distributions are large enough to make mortality comparisons at the top of the white male distribution infeasible. For example, the black male survivor function does not generate a median age of death of 80.3, the median age of death for median-earner white

Figure 5. Survivor Rates for White Male Retirees Born in 1920, Flat Earnings Profile
(In percent)



Note: This figure displays predicted survivor rates for white male retired workers with flat earnings profiles born in 1920 at the 10th, 50th and 90th percentile of the distribution of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars) for all white male retirees. The survivor rates were predicted using the logit models in Table 2, with all other variables set at their cohort means.

males, until about the 96th percentile of the black male earnings distribution. On the other hand, points along the black earnings distribution are feasible for whites, even if incomplete. At the 10th and 90th percentile values of the black male distribution, the median ages of death for white males are 79 and 81.3, respectively, compared with 76.8 and 80.3 for black males. This indicates that the largest differences in mortality between black and white males occur at low earnings levels with the differences narrowing as earnings levels rise. Improvements in mortality at the high end of the white male distribution are outside the range of black male earnings.

An alternative illustration of the black-white income-mortality gap is provided in Figure 7. The figure shows mortality rates by earnings levels for 65-year-old black and white males. The rates decline as earnings levels increase. Differences in the rates decline steadily, particularly at lower earnings levels. The mortality rate difference begins at about 0.7 percentage points at the lowest income level and falls by more than half at about

Table 5. Life Expectancy for Retired Workers Born in 1920

	Flat Earnings Trend				Rising Earnings Trend			
Retirees	10th percentile	Median	90th percentile	90th–10th (months)	10th percentile	Median	90th percentile	90th–10th (months)
Males								
White	78.8	80.3	81.6	33	79.8	81.3	82.6	33
Black	76.8	77.8	80.3	42	77.3	78.3	80.8	42
Female nonduals								
White	82.1	83.1	84.3	26	83.5	84.5	85.8	27
Black	81.1	82.5	84.8	45	82.3	83.8	86.1	45
Female duals								
White								
Own earnings	84.1	85.0	86.4	28	85.1	86.1	87.5	29
Husband's earnings	84.6	85.0	85.4	10	85.7	86.1	86.4	9
Black								
Own earnings	82.1	83.3	85.9	46	83.0	84.3	86.9	47
Husband's earnings	83.2	83.3	83.5	4	84.2	84.3	84.5	4

Source: Authors' calculations using matched records from the Continuous Work History Sample (CWHS) and the Master Beneficiary Records (MBR) of the Social Security Administration.

Note: This table presents the predicted median ages of death, conditional on reaching age 62, for retirees born in 1920. The predictions are based on the logit equations in Tables 2 and 4 and correspond to the 10th, 50th, and 90th percentiles of the group-specific distribution of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars). The workers are classified by sex, by earnings trend and, for females, whether or not they are entitled to benefits under both their own and their spouses' work history (duals and nonduals). The differences between the age of death at the 10th and 90th percentiles are also presented.

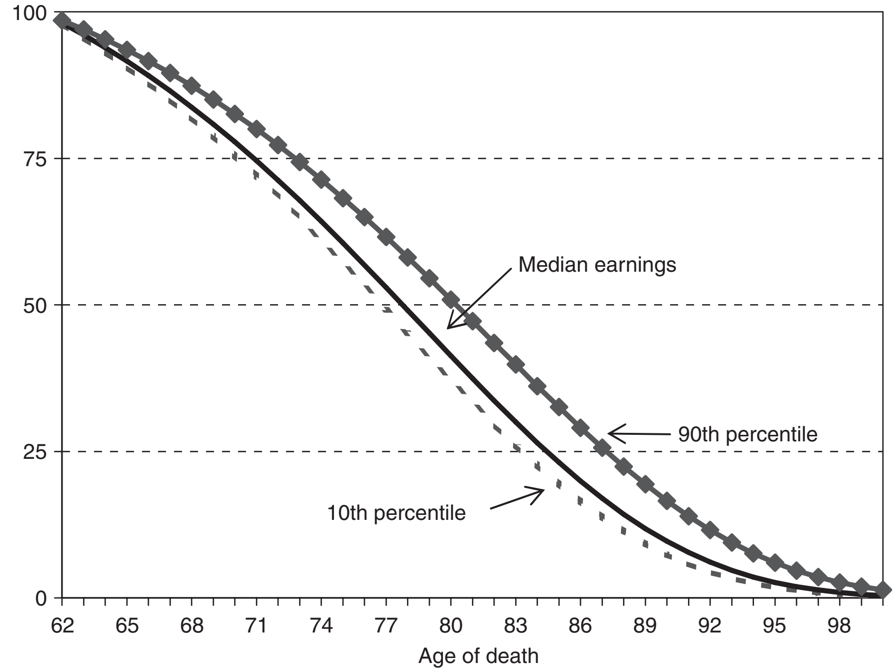
$1.4 million in lifetime earnings, which occurs at the 95th percentile of the black male earnings distribution.

Black and White Females

We define hypothetical female workers in a similar way to that used for males. For nonduals, survivor profiles are shown in Figures 8 and 9 for white and black females, respectively. For white nonduals, the median age of death ranges from 82.1 to 84.3 between the 10th and 90th percentiles of their lifetime earnings distribution, a difference of 26 months (Table 5). For black nonduals, the difference in the median age of death is larger—45 months. For black and white nonduals, a positive earnings trend adds over a year to life expectancy.

For duals there are two sources of lifetime income variation—own and spouses—and we construct survivor profiles by allowing one or the other to

Figure 6. Survivor Rates for Black Male Retirees Born in 1920, Flat Earnings Profile
(In percent)



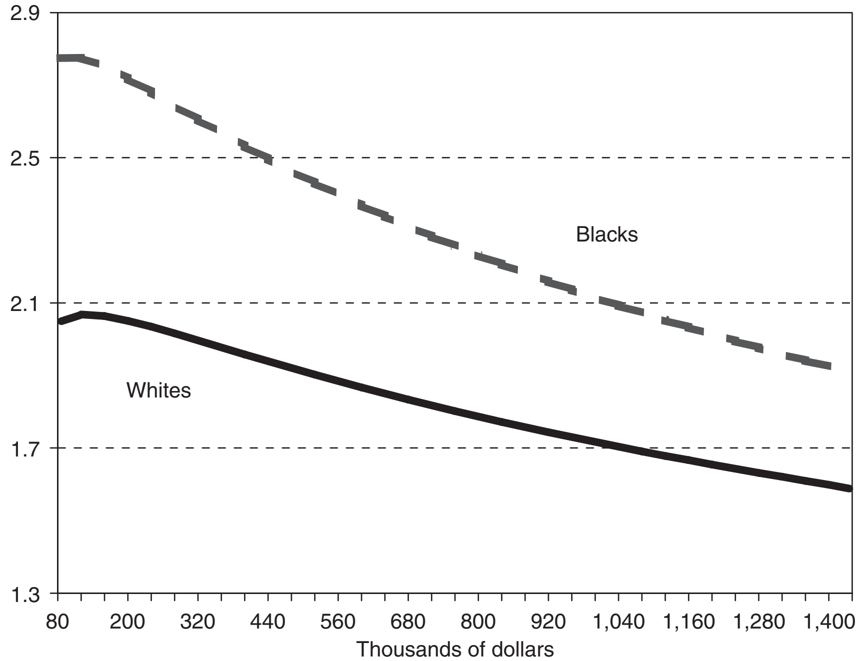
Note: This figure displays predicted survivor rates for black male retired workers with flat earnings profiles born in 1920 at the 10th, 50th, and 90th percentile of the distribution of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars) for all black male retirees. The survivor rates were predicted using the logit model in Table 2, with all other variables set at the cohort means.

vary while holding one to its median value.[14] The results are summarized in Table 5. Higher own lifetime earnings have a significantly negative effect on the mortality of female duals, comparable to that for female nonduals. Higher spousal lifetime earnings has a significantly negative effect for white female duals but an insignificant effect for black duals. For both white and black duals, a positive earnings profile adds about another year of life.

For all groups shown in Table 5, lifetime income has a relatively large positive effect on life expectancy and the effects are remarkably consistent across the groups when evaluating the variation in group-specific income—that is, reading results across each row of the table. In order to make direct comparisons across major demographic groups we computed predicted

[14]Allowing both to vary at the same time would provide the largest income-related differences in life expectancy so our simulations are conservative in this respect. Because we found little empirical relationship between own and spouse's lifetime earnings we allow the income variables to vary independently.

Figure 7. Mortality Rates for 65-Year Old Men, by Lifetime Earnings Levels
(In percent)



Source: Authors' calculations using matched records from the CWHS and MBR of the SSA.

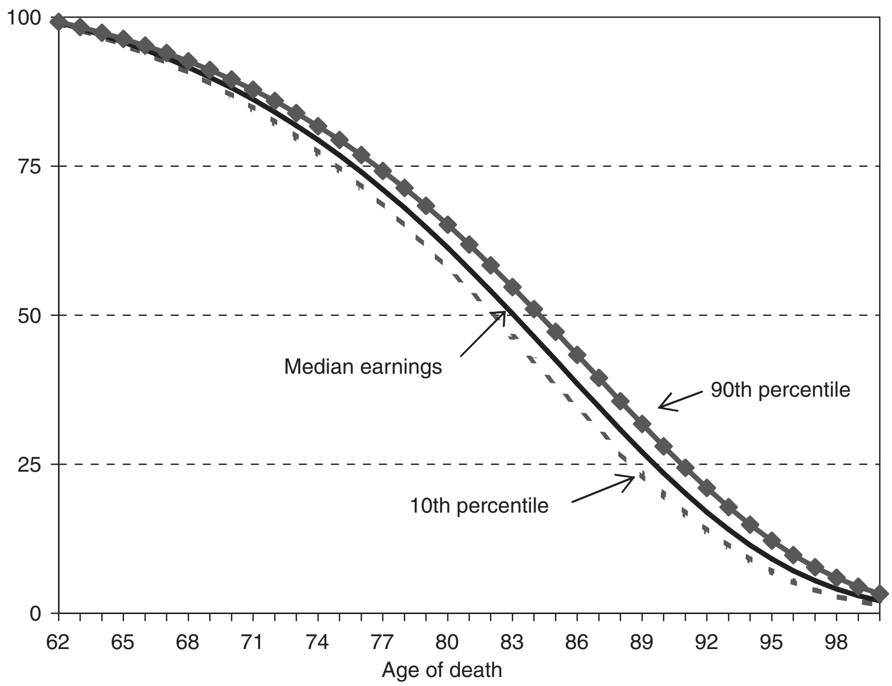
Note: This figure displays the relationship between predicted mortality rates for black and white 65-year old males and lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars). The mortality rates were predicted using the logit model in Table 2, with all other variables set at their cohort means.

median ages of death for white and black males and white and black nondual females at two points from the earnings distributions represented in Figure 4: the fifth percentile from the white male distribution and the 95th percentile from the nondual black female distribution.[15] The results are shown in Table 6.

The difference in age of death between white and black males is about 26 months at the low-income level but this difference narrows substantially to 15 months at the higher-income level. The age-of-death difference between white males and females is about three years at the low-income level and this difference widens somewhat at the higher-income level. The differences between black males and black females are substantial at both income levels

[15]The two lifetime income values are $77,642 and $1,163,370 and correspond to the following percentiles from the respective earnings distributions: white males—5th and 65th; black males—12th and 88th; white females—9th and 95th; black females—17th and 95th. The two points are not feasible for own income of duals.

Figure 8. Survivor Rates for White Female Nonduals Born in 1920, Flat Earnings Profile
(In percent)



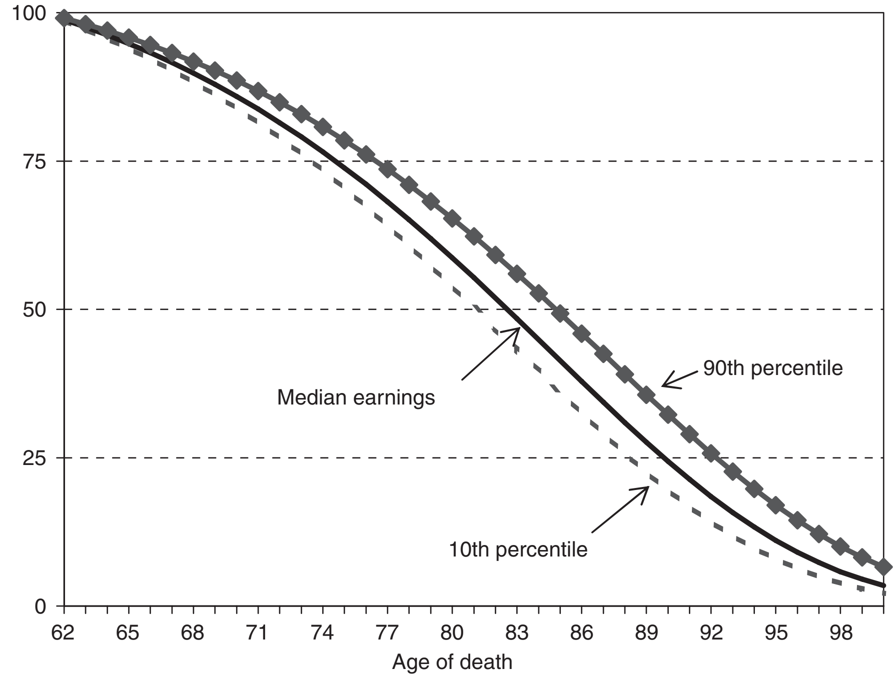
Note: This figure displays predicted survivor rates for white female retired workers with flat earnings profiles born in 1920 at the 10th, 50th, and 90th percentile of the distribution of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars) for all white female retirees who are not entitled to benefits under their spouses' work histories. The survivor rates were predicted using the logit models in Table 4.

but the age-of-death differences are smaller between the two income levels for white and black females. Apparently, age-of-death differences between whites and blacks are largest at low-income levels but the differences between males and females are large and persistent across income levels.

V. Conclusion

This paper provides strong evidence, based on a large file of current and former workers, that mortality is negatively related to lifetime income. For black and white males and females the difference in age of death between low and high lifetime income is on the order of two to three years. Workers with positively trended earnings over their work life may live an additional six to 18 months. Income-related mortality differences between blacks and whites are largest at low-income levels, particularly for males, and narrow

Figure 9. Survivor Rates for Black Female Nonduals Born in1920, Flat Earnings Profile
(In percent)



Note: This figure displays predicted survivor rates for black female retired workers with flat earnings profiles born in 1920 at the 10th, 50th, and 90th percentile of the distribution of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars) for all black female retirees who are not entitled to benefits under their spouses' work histories. The survivor rates were predicted using the logit models in Table 4.

substantially at higher income levels. On the other hand, gender differences in mortality appear to be large and persistent across income levels.

Some might argue that our conclusions are flawed due to the presence of a reverse causation, namely from higher mortality rates to lower incomes. Note, however, that the ages used to construct our lifetime income variable are prior to age 61, and therefore are separate from the years we examine in our mortality function estimates. As a consequence, an early age of death cannot affect the lifetime income of an individual in our logit regression data set. Moreover, we separate out disabled beneficiaries, thereby minimizing the possibility that people with dramatically lower life expectancies have lower measured lifetime incomes.

It remains a possibility that it is not income per se that makes individuals with higher incomes live longer, that instead there may be an underlying demographic or other factor at work. For example, some people who are intelligent, educated, or diligent enough to take good care of their health may

Table 6. Differences Among Median Ages of Death

Category of retiree	Low Income ($77,642) Age of death	Low Income ($77,642) Differences (in months)	High Income ($1,163,370) Age of death	High Income ($1,163,370) Differences (in months)
White males	78.9		80.8	
Less:				
Black male		26		15
White female (nonduals)		−38		−47
Black males	76.8		79.5	
Less:				
Black females (nonduals)		−54		−73
White females (nondual)	82.1		84.7	
Less:				
Black females (nonduals)		10		−11
Black females (nondual)	81.3		85.6	

Source: Authors' calculations using matched records from the Continuous Work History Sample (CWHS) and the Master Beneficiary Records (MBR) of the Social Security Administration.

Note: This table presents differences among the predicted median ages of death, conditional on reaching age 62, for retirees classified by sex and race at low (the 5th percentile of the distribution for white males) and high (the 95th percentile of the distribution for nondual black females) levels of lifetime earnings (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars). Only females who are not entitled to benefits under their spouses' work history are included. All other variables are set to subgroup means.

also attain high incomes because of those same traits. Or, some disadvantaged groups may have low incomes and also lower life expectancies because of childhood-nutrition issues. We have two defenses against this criticism, however. First, we are not trying to conclude that high-income individuals purchase higher life expectancy, and we accept the possibility that some latent variable leads to the positive relationship we observe. Second, our analysis of the income-mortality nexus for dually entitled females shows that the income of the *spouse* (the primary earner) has a significant effect on mortality for white females.

The income-related differences in life expectancy are substantial enough to require consideration when evaluating the distributional consequences of proposals to modify various features of the social security program or when evaluating the existing program. Income-related differences in life expectancy curtail, but do not reverse, the progressivity of the basic social security benefit structure (Liebman, 2002) and have significant implications for individual-account proposals (Brown, 2002; Liebman, 2002).[16] In fact, the income-mortality relationship has implications for a wide range of social security program rules. For example, many reform proposals include a provision to raise the normal retirement age or to increase the early-retirement age. Analyses of such proposals that do not account for the effect of income on life expectancy will misrepresent the differential effects of these proposals on those persons who currently retire early compared with those who retire at later ages.

From a somewhat different perspective, the feedback of life expectancy on retirement decisions complicates any attempt to adjust benefits by age of retirement. The "actuarially fair" adjustment—defined as the adjustment that leaves the internal rate of return on contributions unchanged—will vary across income groups (as well as other groups with systematic differences in life expectancy). We can expect that for any given adjustment, retirees are likely to choose a retirement age that benefits them. Unless the adjustment is either extremely high or extremely low—leading to corner solutions—low-income workers will retire at an earlier age than high-income workers, and both will gain from their choices. A single adjustment factor will cause workers to select "against" the system.[17]

APPENDIX I. The Role of Income Variables in the Mortality Model

Income variables as a group (including the lifetime income and the income trend variables) play two important roles in our mortality model. Most important, they demonstrate the sensitivity of mortality risk to lifetime income. To show this, we estimated the mortality model with and without income variables. Table A1 presents predicted life expectancies—by type of beneficiary and conditional on reaching age 62—of beneficiaries born between 1920 and 1924. The table presents life expectancy estimates corresponding to the mean values of our explanatory variables for individuals in the bottom and top deciles of the distribution of lifetime income, evaluated using the models with and without the income variables. Absent income variables, the estimated life expectancies are virtually identical. When the income variables are included, the minimum differential in life expectancy across these two deciles is 2.8 years for white nondual females, ranging up to 5.8 years for black dual females.

The second role of the income variables is to alter the estimated relationship between mortality and birth cohort. Because lifetime income and birth cohort are correlated, excluding lifetime income from the model biases the coefficient on birth cohort. For instance, for white males, the estimated coefficient on birth cohort is -0.0173 when the income variables are included and -0.0194 when the income variables are excluded. As a result, the model without income variables incorrectly attributes the increase in life expectancy over time due to the trend in lifetime income to the passage of time, which, in

[16]See also Duggan, Gillingham, and Greenlees (1993 and 1996) on the distributional impacts of social security.

[17]The implications of constant actuarial adjustments across cohorts and income classes are analyzed in Duggan and Soares (2002).

Table A1. Life Expectancy Predicted With and Without Income Variables

	Birth Years 1920–24					
	Full model			No income variables		
Retirees	Bottom decile	Top decile	Difference	Bottom decile	Top decile	Difference
Males						
White	79.8	83.2	3.4	80.8	81.0	0.2
Black	77.5	80.7	3.2	78.1	78.2	0.1
Nondual females						
White	79.8	82.7	2.8	80.8	80.8	0.1
Black	78.2	82.3	4.1	79.6	79.6	0.0
Dual females						
White	81.7	84.8	3.2	82.7	82.7	0.0
Black	79.2	84.9	5.8	81.1	81.1	0.0

Note: This table presents predicted median life expectancies for workers with birth years 1920 to 1924 classified by race, by sex, and, for females, according to whether the worker was entitled to a benefit under her spouse's work history. The predictions are calculated at the mean levels of explanatory variables for workers in the bottom and top deciles of the distribution of lifetime income (aggregate labor earnings between ages 35 and 60, measured in 2005 dollars) using both the full logit model and a logit model that excludes all income variables.

turn, would bias projections of future cash flows in pension programs. In fact, the bias is so great that the estimated difference in life expectancy between the top and bottom deciles of the entire sample of white males is greater when we exclude the income variables than when we include them.

APPENDIX II. Summary Statistics for Logit Models

See Tables B1 and B2.

Table B1. Summary Statistics of Variables in Tables 2 and 4

Variables	Means	Std. dev.	Variances and Covariances										
			Hazard	Age	Birth year	LE	LE*2	LE*3	Rising trend	Flat trend	LESP	LESP*2	LESP*3
Table 2—White males													
N = 3,841,182													
Hazard	0.039	0.194	0.038	0.216	−0.159	−0.007	−0.191	−3.813	0.001	−0.001			
Age	70.617	6.782	0.216	46.001	−22.051	−0.616	−14.925	−288.302	0.455	−0.111			
Birth year	17.755	10.390	−0.159	−22.051	107.943	3.746	95.542	1,873.712	−1.636	0.431			
LE	13.295	1.518	−0.007	−0.616	3.746	2.303	43.513	733.476	0.038	0.004			
LE*2	179.066	29.994	−0.191	−14.925	95.542	43.513	899.630	15,853.099	0.372	0.956			
LE*3	2,424.230	534.256	−3.813	−288.302	1,873.712	733.476	15,853.099	285,429.165	3.522	24.166			
Rising trend	0.539	0.498	0.001	0.455	−1.636	0.038	0.372	3.522	0.249	−0.118			
Flat trend	0.218	0.413	−0.001	−0.111	0.431	0.004	0.956	24.166	−0.118	0.171			
Table 2—Black males													
N = 310,571; Mean of dependent variable = 0.046													
Hazard	0.046	0.209	0.044	0.218	−0.142	−0.008	−0.217	−4.168	0.001	−0.001			
Age	70.217	6.683	0.218	44.668	−23.076	−0.920	−20.947	−381.357	0.420	−0.148			
Birth year	17.670	10.704	−0.142	−23.076	114.582	4.939	118.428	2,203.371	−1.547	0.682			
LE	12.592	1.664	−0.008	−0.920	4.939	2.769	51.434	841.675	0.085	−0.001			
LE*2	161.316	32.289	−0.217	−20.947	118.428	51.434	1,042.586	17,792.696	1.355	1.025			
LE*3	2,082.660	556.607	−4.168	−381.357	2,203.371	841.675	17,792.696	309,811.683	19.845	26.244			
Rising trend	0.552	0.497	0.001	0.420	−1.547	0.085	1.355	19.845	0.247	−0.100			
Flat trend	0.181	0.385	−0.001	−0.148	0.682	−0.001	1.025	26.244	−0.100	0.148			
Table 4—White female duals													
N = 1,532,488													
Hazard	0.037	0.189	0.036	0.300	−0.153	−0.006	−0.141	−2.518	−0.001	0.000	−0.006	−0.151	−3.014
Age	72.146	7.525	0.300	56.632	−26.740	−0.162	−5.082	−98.833	0.157	−0.043	−0.894	−23.266	−463.484
Birth year	17.671	9.424	−0.153	−26.740	88.819	1.063	32.420	623.753	−0.191	0.159	3.306	84.615	1,669.229
LE	11.418	2.338	−0.006	−0.162	1.063	5.466	78.827	1,091.088	0.264	−0.313	0.084	1.265	15.631
LE*2	135.845	35.335	−0.141	−5.082	32.420	78.827	1,248.582	18,318.633	3.792	−3.563	2.695	54.061	897.607
LE*3	1,629.960	527.267	−2.518	−98.833	623.753	1,091.088	18,318.633	278,010.665	52.440	−39.572	52.154	1,111.105	19,276.322
Rising trend	0.555	0.497	−0.001	0.157	−0.191	0.264	3.792	52.440	0.247	−0.069	−0.018	−0.455	−8.918
Flat trend	0.124	0.329	0.000	−0.043	0.159	−0.313	−3.563	−39.572	−0.069	0.109	0.012	0.301	5.883
LESP	13.445	1.277	−0.006	−0.894	3.306	0.084	2.695	52.154	−0.018	0.012	1.631	30.630	518.139

Table B1 *(concluded)*

LESP*2	182.400	25.150	−0.151	−23.266	84.615	1.265	54.061	1,111.105	−0.455	0.301	30.630	632.530	11,221.257
LESP*3	2,483.010	451.291	−3.014	−463.484	1,669.229	15.631	897.607	19,276.322	−8.918	5.883	518.139	11,221.257	203,663.617
Table 4—White female nonduals													
N = *1,608,186*													
Hazard	0.046	0.209	0.044	0.323	−0.229	−0.011	−0.254	−4.678	0.001	−0.001			
Age	71.558	7.577	0.323	57.412	−32.882	−1.156	−25.299	−452.368	0.365	−0.140			
Birth year	17.051	10.992	−0.229	−32.882	120.832	4.792	109.048	1,974.674	−1.028	0.474			
LE	12.469	1.731	−0.011	−1.156	4.792	2.995	49.971	770.818	0.112	−0.080			
LE*2	158.475	30.320	−0.254	−25.299	109.048	49.971	919.301	14,948.785	1.454	−0.384			
LE*3	2,026.010	499.760	−4.678	−452.368	1,974.674	770.818	14,948.785	249,759.817	18.393	3.131			
Rising trend	0.681	0.466	0.001	0.365	−1.028	0.112	1.454	18.393	0.217	−0.096			
Flat trend	0.141	0.348	−0.001	−0.140	0.474	−0.080	−0.384	3.131	−0.096	0.121			
Table 4—Black female duals													
N = *86,385*													
Hazard	0.038	0.191	0.036	0.264	−0.128	−0.012	−0.265	−4.566	0.000	−0.001	−0.004	−0.120	−2.388
Age	71.550	7.266	0.264	52.796	−26.113	−1.251	−28.695	−498.181	0.289	−0.094	−0.977	−24.393	−461.980
Birth year	15.478	8.857	−0.128	−26.113	78.443	4.498	100.627	1,727.092	−0.580	0.254	3.012	75.930	1,440.474
LE	11.452	1.434	−0.012	−1.251	4.498	2.056	37.247	582.406	0.076	−0.032	0.481	10.975	197.592
LE*2	133.210	27.113	−0.265	−28.695	100.627	37.247	735.099	11,956.575	1.246	0.050	10.398	241.412	4,386.426
LE*3	1,562.800	445.058	−4.566	−498.181	1,727.092	582.406	11,956.575	198,076.802	18.968	5.917	176.424	4,124.119	75,233.265
Rising trend	0.619	0.486	0.000	0.289	−0.580	0.076	1.246	18.968	0.236	−0.061	−0.066	−1.631	−30.430
Flat trend	0.098	0.297	−0.001	−0.094	0.254	−0.032	0.050	5.917	−0.061	0.088	0.014	0.346	6.515
LESP	12.672	1.555	−0.004	−0.977	3.012	0.481	10.398	176.424	−0.066	0.014	2.418	43.125	694.377
LESP*2	162.990	29.099	−0.120	−24.393	75.930	10.975	241.412	4,124.119	−1.631	0.346	43.125	846.766	14,310.548
LESP*3	2,108.480	497.592	−2.388	−461.980	1,440.474	197.592	4,386.426	75,233.265	−30.430	6.515	694.377	14,310.548	247,597.503
Table 4—Black female nonduals													
N = *190,498*													
Hazard	0.047	0.211	0.045	0.285	−0.219	−0.017	−0.397	−7.159	0.001	−0.001			
Age	71.197	7.498	0.285	56.213	−33.834	−2.391	−50.397	−872.385	0.414	−0.192			
Birth year	17.916	11.111	−0.219	−33.834	123.465	9.018	196.758	3,434.903	−1.150	0.627			
LE	11.977	1.875	−0.017	−2.391	9.018	3.514	59.873	925.491	0.091	−0.070			
LE*2	146.959	33.558	−0.397	−50.397	196.758	59.873	1,126.128	18,301.930	0.842	0.029			
LE*3	1,819.980	552.096	−7.159	−872.385	3,434.903	925.491	18,301.930	304,809.602	6.916	11.466			
Rising trend	0.690	0.463	0.001	0.414	−1.150	0.091	0.842	6.916	0.214	−0.092			
Flat trend	0.134	0.340	−0.001	−0.192	0.627	−0.070	0.029	11.466	−0.092	0.116			

Table B2. Summary Statistics of Parameters in Tables 2 and 4

Variables	Means	Variances and Covariances (value × 1,000,000)										
		Intercept	Age	Birth year	LE	LE*2	LE*3	Rising trend	Flat trend	LESP	LESP*2	LESP*3
Table 2—White males												
N = 3,841,182; mean of dependent variable = 0.039												
Intercept		2,084	−10.000	−5.230	−270.000	20.000	−0.433	−40.000	−80.000			
Age	70.617	−10	0.131	0.040	−0.046	0.014	−0.001	−0.130	−0.038			
Birth year	17.755	−5	0.040	0.125	−1.340	0.250	−0.011	0.812	0.636			
LE	13.295	−270	−0.046	−1.340	534.000	−80.000	3.055	−10.000	−4.980			
LE*2	179.066	20	0.014	0.250	−80.000	12.000	−0.479	2.305	2.268			
LE*3	2,424.230	0	−0.001	−0.011	3.055	−0.479	0.019	−0.108	−0.125			
Rising trend	0.539	−40	−0.130	0.812	−10.000	2.305	−0.108	51.000	38.000			
Flat trend	0.218	−80	−0.038	0.636	−4.980	2.268	−0.125	38.000	77.000			
Table 2—Black males												
N = 310,571; mean of dependent variable = 0.046												
Intercept		20,192	−110.000	−50.000	−2,610.000	197.000	−4.610	−310.000	−870.000			
Age	70.217	−110	1.417	0.447	−1.250	0.261	−0.012	−1.320	−0.365			
Birth year	17.670	−50	0.447	1.170	−8.110	1.611	−0.078	6.676	3.166			
LE	12.592	−2,610	−1.250	−8.110	4,054.000	−610.000	24.000	−40.000	15.000			
LE*2	161.316	197	0.261	1.611	−610.000	97.000	−3.940	12.000	16.000			
LE*3	2,082.660	−5	−0.012	−0.078	24.000	−3.940	0.162	−0.694	−1.060			
Rising trend	0.552	−310	−1.320	6.676	−40.000	12.000	−0.694	474.000	329.000			
Flat trend	0.181	−870	−0.365	3.166	15.000	16.000	−1.060	329.000	863.000			
Table 4—White female duals												
N = 1,532,488; mean of dependent variable = 0.037												
Intercept		7,542	−30.000	−20.000	−150.000	4.460	0.208	−40.000	−270.000	−970.000	85.000	−2.390
Age	72.146	−30	0.399	0.208	−0.081	0.028	−0.002	−0.316	−0.123	0.154	−0.023	0.001
Birth year	17.671	−20	0.208	0.468	−0.696	0.174	−0.009	0.010	−0.136	−5.200	0.911	−0.040
LE	11.418	−150	−0.081	−0.696	785.000	−130.000	5.730	−10.000	−30.000	−40.000	7.393	−0.305
LE*2	135.845	4	0.028	0.174	−130.000	23.000	−1.020	3.278	11.000	6.960	−1.230	0.052
LE*3	1,629.960	0	−0.002	−0.009	5.730	−1.020	0.045	−0.170	−0.569	−0.245	0.046	−0.002
Rising trend	0.555	−40	−0.316	0.010	−10.000	3.278	−0.170	95.000	59.000	1.447	−0.425	0.026
Flat trend	0.124	−270	−0.123	−0.136	−30.000	11.000	−0.569	59.000	291.000	13.000	−2.400	0.104
LESP	13.445	−970	0.154	−5.200	−40.000	6.960	−0.245	1.447	13.000	2,417.000	−360.000	14.000

Table B2 *(concluded)*

LESP*2	182.400	85	−0.023	0.911	7.393	−1.230	0.046	−0.425	−2.400	−360.000	55.000	−2.130
LESP*3	2,483.010	−2	0.001	−0.040	−0.305	0.052	−0.002	0.026	0.104	14.000	−2.130	0.083
Table 4—White female nonduals												
N = *1,608,186; Mean of dependent variable* = *0.046*												
Intercept		3,142	−20.000	−10.000	−310.000	22.000	−0.433	−80.000	−230.000			
Age	71.558	−20	0.251	0.104	−0.030	0.019	−0.001	−0.231	−0.055			
Birth year	17.051	−10	0.104	0.261	−1.670	0.333	−0.016	1.008	0.483			
LE	12.469	−310	−0.030	−1.670	868.000	−140.000	5.642	−3.660	7.460			
LE*2	158.475	22	0.019	0.333	−140.000	23.000	−0.941	1.537	3.139			
LE*3	2,026.010	0	−0.001	−0.016	5.642	−0.941	0.039	−0.101	−0.225			
Rising trend	0.681	−80	−0.231	1.008	−3.660	1.537	−0.101	112.000	90.000			
Flat trend	0.141	−230	−0.055	0.483	7.460	3.139	−0.225	90.000	231.000			
Table 4—Black female duals												
N = *86,385; Mean of dependent variable* = *0.038*												
Intercept		156,787	−610.000	−420.000	−13,290.000	1,195.000	−30.000	−970.000	−5,560.000	−16,730.000	1,581.000	−50.000
Age	71.550	−610	7.250	4.283	−0.126	−0.104	0.003	−4.570	−2.430	0.886	0.007	−0.003
Birth year	15.478	−420	4.283	9.332	−0.402	0.970	−0.104	9.029	−1.920	−40.000	7.443	−0.343
LE	11.452	−13,290	−0.126	−0.402	20,590.000	−3,340.000	143.000	−470.000	184.000	−620.000	116.000	−5.190
LE*2	133.210	1,195	−0.104	0.970	−3,340.000	569.000	−20.000	99.000	96.000	148.000	−30.000	1.253
LE*3	1,562.800	−30	0.003	−0.104	143.000	−20.000	1.095	−5.050	−6.720	−7.080	1.354	−0.062
Rising trend	0.619	−970	−4.570	9.029	−470.000	99.000	−5.050	1,776.000	1,216.000	50.000	−20.000	1.467
Flat trend	0.098	−5,560	−2.430	−1.920	184.000	96.000	−6.720	1,216.000	5,754.000	266.000	−50.000	2.585
LESP	12.672	−16,730	0.886	−40.000	−620.000	148.000	−7.080	50.000	266.000	23,916.000	−3,600.000	142.000
LESP*2	162.990	1,581	0.007	7.443	116.000	−30.000	1.354	−20.000	−50.000	−3,600.000	566.000	−20.000
LESP*3	2,108.480	−50	−0.003	−0.343	−5.190	1.253	−0.062	1.467	2.585	142.000	−20.000	0.926
Table 4—Black female nonduals												
N = *190,498; Mean of dependent variable* = *0.047*												
Intercept		24,736	−170.000	−90.000	−2,570.000	196.000	−4.350	−610.000	−2,060.000			
Age	71.197	−170	1.991	0.829	3.152	−0.409	0.015	−2.010	−0.597			
Birth year	17.916	−90	0.829	2.308	−10.000	2.437	−0.134	10.000	4.552			
LE	11.977	−2,570	3.152	−10.000	5,233.000	−850.000	35.000	−190.000	−3.620			
LE*2	146.959	196	−0.409	2.437	−850.000	142.000	−6.070	39.000	41.000			
LE*3	1,819.980	−4	0.015	−0.134	35.000	−6.070	0.262	−1.930	−2.620			
Rising trend	0.690	−610	−2.010	10.000	−190.000	39.000	−1.930	888.000	704.000			
Flat trend	0.134	−2,060	−0.597	4.552	−3.620	41.000	−2.620	704.000	2,080.000			

REFERENCES

Allison, P., 1995, *Survival Analysis Using the SAS System, A Practical Guide* (Cary, North Carolina: SAS Press).

Aziz, F., and W. Buckler, 1992, "The Status of Death Information in Social Security Administration Files," paper presented at the Joint Statistical Meetings, Boston, Massachusetts.

Brown, J., 2002, "Differential Mortality and the Value of Individual Account Retirement Annuities," in *The Distributional Aspects of Social Security and Social Security Reform*, ed. by M. Feldstein and J. Liebman (Chicago: University of Chicago Press).

Committee on Finance of the U.S. Senate and the Committee on Ways and Means of the U.S. House of Representative 1976, Report of the Consultant Panel on Social Security, August 1976, U.S. Government Printing Office, Washington DC, No. 70–577.

Congressional Budget Office, 2006, "Is Social Security Progressive?" Economic and Budget Issue Brief (December 15).

Duggan, J.E., 1996, "Distributional Effects of Social Security: The Notch Issue Revisited," *Public Finance Quarterly*, Vol. 24, No. 3 (July), pp. 349–70.

Duggan, J.E., and R. Gillingham, 1999, "The Effect of Errors in the CPI on Social Security Finances," *Journal of Business and Economic Statistics*, Vol. 17, No. 2 (April), pp. 161–9.

Duggan, J.E., and C. Soares, 2002, "Actuarial Nonequivalence in Early and Delayed Social Security Benefit Claims," *Public Finance Review*, Vol. 30, No. 3 (May), pp. 188–207.

Duggan, J.E., R. Gillingham, and J. Greenlees, 1993, "The Returns Paid to Early Social Security Cohorts," *Contemporary Policy Issues*, Vol. 24, No. 3 (October), pp. 1–13.

Efron, B., 1988, "Logistic Regression, Survival Analysis, and the Kaplan-Meier Curve," *Journal of the American Statistical Association*, Vol. 3, No. 402 (June), pp. 414–25.

Friedman, M., 1972, "Second Lecture," in *Social Security: Universal or Selective?* ed. by W. Cohen and M. Friedman (Washington: American Enterprise Institute).

Hadley, J., and A. Osei, 1982, "Does Income Affect Mortality? An Analysis of the Effects of Different Types of Income on Age/Sex/Race-Specific Mortality Rates in the United States," *Medical Care*, Vol. 20, No. 9 (September), pp. 901–14.

House, J., R. Kessler, and A. Herzog, 1990, "Age, Socioeconomic Status, and Health," *The Milbank Quarterly*, Vol. 68, No. 3, pp. 383–411.

Hoyert, D., G. Singh, and H. Rosenberg, 1995, "Sources of Data on Socioeconomic Differential Mortality in the United States," *Journal of Official Statistics*, Vol. 11, No. 3, pp. 233–60.

Kitagawa, E., and P. Hauser, 1973, *Differential Mortality in the U.S.: A Study of Socioeconomic Epidemiology* (Cambridge, Massachusetts: Harvard University Press).

Krueger, P., and others, 2003, "Socioeconomic Status and Age: The Effect of Income Sources and Portfolios on U.S. Adult Mortality," *Sociological Forum*, Vol. 18, No. 3, pp. 465–82.

Liebman, J., 2002, "Redistribution in the Current U.S. Social Security System," in *The Distributional Aspects of Social Security and Social Security Reform*, ed. by M. Feldstein and J. Liebman (Chicago: University of Chicago Press).

Myers, R., 1993, *Social Security* (Philadelphia: University of Pennsylvania Press).

Rosen, S., and P. Taubman, 1979, "Changes in the Impact of Education and Income on Mortality in the U.S." in *Statistics of Income and Related Administrative Record Research* (Washington: Department of Treasury).

Sorlie, P., E. Backlund, and J. Keller, 1995, "U.S. Mortality by Economic, Demographic, and Social Characteristics: The National Longitudinal Mortality Study," *American Journal of Public Health*, Vol. 85, No. 7, pp. 949–56.

Wilmoth, J., and M. Dennis, 2006, "Social Differences in Older Adult Mortality in the United States: Questions, Data, Methods, and Results," in *Human Longevity, Individual Life Duration, and the Growth of the Oldest-Old Population*, ed. by Jean-Marie Robine and others (Oxford, UK: Oxford University Press).

Zayatz, T., 1999, *Social Security Disability Program Worker Experience*, Actuarial Study No. 114 (Washington: Office of the Actuary, Social Security Administration).