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**Progressive Returns to Social Security?  
An Answer from Social Security Records**

by

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## I. Introduction

A basic redistributive goal of the U. S. Social Security program is to award higher rates of return on the contributions of low-wage than high-wage workers. In practice, this goal can be thwarted if low-income and high-income beneficiaries exhibit offsetting characteristics.<sup>1</sup> Over twenty years ago Milton Friedman stated that high income persons clearly receive higher Social Security benefits relative to contributions paid than low-income persons because, *inter alia*, persons in high income classes have a higher life expectancy, and so will tend to receive benefits for a longer period of time (Friedman, 1972, p. 35).

Aaron (1977) furnished evidence that showed differential mortality rates may fully offset the progressivity of the Social Security benefit formula, and he suggested the formula be made more progressive. Leimer (1978) extended Aaron's analysis and concluded differential mortality does *not* upset the progressivity of the Social Security benefit structure.

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<sup>1</sup>The progressive nature of the Social Security benefit structure is achieved through a nonlinear benefit formula, a cap on taxable wages, and taxation of benefits. For example, in 1996 the basic Social Security benefit formula was 90% of the first \$437 of average indexed monthly earnings, 32% of the next \$2,198, and 15% of average indexed monthly earnings over \$2,635.

Over the past two decades, the mortality literature has converged on the finding that life expectancy is higher for higher income persons (Feinstein, 1993) and recent research has found a widening gap in the mortality of high and low income persons (Duleep, 1989; Pappas, *et al*, 1993). Yet, since the Aaron and Leimer papers, during a period that has included major reforms to the Social Security program, little attempt has been made to investigate the implications of this research for benefit progressivity. Recent exceptions are papers by Garrett (1995) and Panis and Lillard (1995). Neither paper examines actual Social Security program experience, however. Panis and Lillard use survey data to estimate the effect of income on mortality and employ a representative worker approach to investigate progressivity. Their findings show a reversal in progressivity for unmarried males. Garrett uses a representative beneficiary along with parameters from prior income-mortality research. He finds a significant offsetting effect on benefit progressivity. The representative individual approach has well-known shortcomings, however, and can give misleading results, a fact that eventually led Aaron to reverse his conclusion about the progressivity of Social Security.<sup>2</sup>

In contrast, this paper investigates the effect of endogenous mortality on the benefit structure of the U.S. Social Security program using a large file of administrative records. Our investigation has two parts. First we analyze the income-mortality relationship in our data by estimating death hazard rate functions that allow for the effects of demographic and economic factors. We estimate these functions for a sample of persons born during the period 1895 to 1922 and covered under the Social Security program. Our focus is on the role of income in determining mortality rates. We find income has a significantly negative effect on the probability of death for men and women, a result consistent with recent mortality research. Next, we incorporate the

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<sup>2</sup>Aaron (1982) concluded the Social Security benefit formula is progressive, pointing to the work of Freiden, Leimer, and Hoffman (1976) that, as in this paper, used longitudinal administrative records. In contrast to this paper, Freiden, Leimer, and Hoffman did not include survivors and did not examine explicitly the income-mortality relationship.

income-adjusted mortality rates into a detailed simulation of lifetime benefits and real rates of return for a subset of the 1895-1922 cohorts. Here we find income-adjusted mortality rates do affect the distribution of Social Security benefits across income classes, but not nearly enough to alter the basic progressivity of the benefit structure.

Our empirical work is based on individual administrative records from the Social Security Administration's (SSA) 1988 Continuous Work History Sample (CWHS). The CWHS is an earnings history sample for one percent of Social Security records. It contains over 2.5 million records with actual earnings histories spanning the period 1951 to 1988. The file contains month and year of death information for deceased persons covered by the Social Security program. The CWHS also has benefit information that allows us to compute complete benefit histories for current and former Social Security beneficiaries. Further discussion of the CWHS is provided later in the paper and in an accompanying appendix.

This paper is the first to base an investigation of the effect on Social Security progressivity of variation in mortality rates by income on actual program experience. Our analysis also advances earlier work in this area by explicitly including survivor benefits. In section II we briefly review previous research on the relationship between income and mortality and present our estimates of this relationship. In section III we analyze the implications of our estimated mortality rates for the distribution of net returns to Social Security beneficiaries. Section IV is a summary and conclusion.

## **II. Income and Mortality**

### *IIa. Previous Research*

In theory, the effect of income on mortality cannot be determined *a priori*. Higher income may mean higher mortality if the consumption of some goods that adversely affect health (rich

foods, for example) increases with income.<sup>3</sup> On the other hand, higher income may result in increased consumption of "healthy" (more expensive) foods and other life-prolonging goods and services. Though empirical evidence is mixed, recent research employing improved data sets generally finds an inverse relation between income and mortality.<sup>4</sup> Using a variety of data bases, Kitagawa and Hauser (1973), Rosen and Taubman (1979), Caldwell and Diamond (1979), Hadley and Osei (1982), Duleep (1986), Rogers (1992), Sorlie, *et al* (1992), and Menchik (1993) all found a strong inverse relation between income and mortality.<sup>5</sup>

Other factors may also affect mortality. A higher level of education suggests increased knowledge about health and health services, though this effect may operate primarily through income (Feinstein, 1992). Recently, several researchers have attempted to assess the comparative roles of race and other socioeconomic variables in determining death hazards. Rogers (1992)

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<sup>3</sup>Fuchs (1974) argued that as income rises there is a tendency for the elasticity of mortality with respect to income to become positive.

<sup>4</sup>Studies that found a direct relation between income and the probability of death generally relied on data characterized by a high level of aggregation, such as SMSAs, states, census tracts, etc. Kitawaga and Hauser (1973) provide a thorough review of early studies. Also, see Hadley and Osei (1982) and Duleep (1986).

<sup>5</sup>The data bases used in the studies are as follows. Kitwaga and Hauser: 1960 Census data matched to 1960 death certificate data; Rosen and Taubman: 1973 Current Population Survey-Internal Revenue Service-SSA Exact Match File; Caldwell and Diamond: 1969 IRS data and 1970-1976 SSA data; Hadley and Osie: 1970 Census; Duleep: 1973 Exact Match file with Social Security death data through 1978; Rogers: 1986 National Health Interview Survey and the 1986 National Mortality Followback Survey; Sorlie: National Longitudinal Mortality Study; Menchik: National Longitudinal Survey of Mature Men.

estimated mortality functions, controlling sequentially for demographic, familial status, and income variables. Age, race and sex were found to be significant determinants of mortality. When marital status and family size variables were added to the equation the importance of race was reduced, and when family income was included, the effect of race was eliminated. Menchik (1993) also found the effect of race on mortality eroded once income and family background variables were accounted for. In addition, Menchik found no separate effect of education on mortality.

### *Iib. The Continuous Work History Sample*

For the purpose of mortality research, the CWHS has several advantages over other data sets. It is much larger than most of the available surveys with information on date of death. The CWHS has individual longitudinal information on key variables (e.g., earned income), demographic information (age, race, sex), employment and health status information, all of which are matched to mortality data.<sup>6</sup> This allows us to avoid some of the problems associated with aggregation and one-time measurement of variables of primary interest. For example, previous studies have often categorized people into income classes based on a single year of income. This procedure tends to mix populations that are either temporarily or permanently in a certain income class, obscuring the effects of permanent and transitory income.

The CWHS sample constitutes a very large base of information on the mortality experience of older workers and retirees. The sample also provides us with a large group of newly retired workers whose post-1988 benefit experience can be projected and compared under differing sets

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<sup>6</sup>The quality of death information in the CWHS is summarized in Aziz and Buckler (1992). Generally, the death reporting has improved over time. As could be expected, the information is most reliable for former beneficiaries.

of estimated mortality parameters. The CWHS also has limitations for estimating mortality models. Most important for our purposes is that individual covered earnings is the only source of income data. Also, demographic data are limited to age, sex, and race, so the possible effects of other family background characteristics on mortality rates cannot be evaluated.

The 1988 CWHS has just under 3.4 million records, of which over 2.5 million are "active" accounts, that is, with any history of Social Security earnings. For computational purposes, we drew a ten percent random sample from the 2.5 million records on the full active file. In estimating the income-mortality relationship we then chose to work with the subsample of persons born during the period 1895 to 1923. After editing out anomalous cases, we arrived at a data set with 44,252 records. We confined our attention to mortality probabilities at age 65 or above, which yielded a sample of 205,549 male and 160,009 female person-years covering ages 65 through 93 and years 1960 through 1988.

Our cohort selection reflects the twin requirements of information on lifetime income and information on mortality. Cohorts born prior to 1895 reached age 42 prior to the establishment of Social Security in 1937, so our data on covered earnings provide a relatively poor indication of their financial resources. For cohorts born after 1923, the 1988 CWHS provides no indication of their mortality experience after age 65.

### *IIIc. Estimation of Mortality Models*

Following prior research, we hypothesize a function that relates mortality probability  $M$  to age  $A$ , income  $Y$ , and a vector of demographic variables  $Z$ :

$$(1) \quad M = f(A, Y, Z).$$

Our unit of observation is an individual year of a participant's life, and  $M$  is the discrete "hazard rate," the probability that death will take place in a given year, given the individual was alive at the end of the previous year. We estimate the hazard directly, using the well-known logit probability model:

$$(2) \quad M = 1/(1 + \exp(-bX)),$$

where  $X$  is a vector of explanatory variables including  $A$ ,  $Z$ , and a set of income variables described below. In our logit formulation the coefficients in  $b$  can be interpreted as the effect of a one-unit increase in the associated variable on the logarithm of the ratio  $M/(1-M)$ . Thus, variables with positive logit coefficients are associated with higher death rates and shorter expected remaining lifetimes. We should expect a negative coefficient on age, for example, since the probability of survival declines with age.<sup>7</sup>

The variables in  $Z$  include year of birth, two race variables -- black and other nonwhite (primarily Asians) -- and disability status, measured as receipt of Social Security disability benefits prior to age 65. Income is measured as the log of average covered wages from age 46 to age 60

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<sup>7</sup>We examined the plausibility of the CWHS death data and our logit formulation by comparing estimates of equation (2) with estimates based on mortality data from the Bureau of Census. Using the CWHS, we estimated equation (2) for each sex with age, birth year, and race as explanatory variables. Then, using the Census data, we fit (for males and females separately) semi-logarithmic regressions of the central death rate, for age groups ranging from 45-49 to over 100 and the years 1990, 2000, and 2050, on age, year and dummy variables for black and white race. Predictions from the logit-CWHS model and the Census-based model were extremely close with correlations exceeding .99. We also needed the Census-based model to project mortality rates of spouses, widow, and widowers. Further information is provided in the appendix.

(including wage-and-salary and self-employment income), with each year's earnings measured in thousands of 1988 dollars using the Bureau of Labor Statistics' CPI-U-X1 series for deflation.

In addition, we included two other income-related variables necessitated by the nature of our data set. Some individuals (2.7 percent of males and 6.6 percent of females) had zero taxable earnings throughout ages 46 to 60. Especially for members of the earlier birth cohorts, a year without taxable earnings may reflect work in an uncovered industry or occupation, rather than a truly low level of income. Zero earnings, particularly for women, could also reflect work as a homemaker, and thereby underestimate total household earnings in our data. In order to account for these possibilities, we included a discrete variable indicating zero earnings over ages forty-six to sixty. Our expectation is for negative coefficients on the zero-earnings variable.<sup>8</sup>

Furthermore, in each year earned income is observed in the CWHS only up to the taxable wage ceiling (cap) level, which varied (in 1988 dollars) from \$22,214 in 1941 to \$45,000 in 1988. We therefore included a variable representing the number of "capped" earnings values included in the fifteen-year average. The coefficient should reflect our underestimate of total income in those years.

Table 1 displays our parameter estimates of equation (2), estimated separately for men and women, and either including or excluding income variables. For the most part the coefficients are statistically significant, as would be expected given the sample sizes involved, and the signs of the coefficients are reasonable. For both men and women, the probability of death rises with age and falls with cohort birth year. An additional year of age increases the log-odds of mortality by .076

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<sup>8</sup>The zero income dummy variable also accommodates the log-income variable which we set to zero for individuals with no earnings over ages 46-60.

**Table 1**  
**Estimated Mortality Hazard Models**

<b>Variable</b>	<b><i>Including Income Variables</i></b>		<b><i>Excluding Income Variables</i></b>	
	<b>Men</b>	<b>Women</b>	<b>Men</b>	<b>Women</b>
<i>Intercept</i>	-8.1261 (.1638)	-9.3247 (.2142)	-8.3940 (.1400)	-10.0188 (.2024)
<i>Age</i>	0.0756 (.0018)	0.0907 (.0026)	0.0752 (.0018)	0.0892 (.0026)
<i>Birth Year</i>	-0.0114 (.0017)	-0.0050 <sup>1</sup> (.0027)	-0.0143 (.0017)	-0.0101 (.0026)
<i>log(Income)</i>	-0.0274 (.0105)	-0.1063 (.0106)		
<i>Zero Income</i>	-0.1325 <sup>1</sup> (.1062)	-0.4813 (.1012)		
<i>Cap Income Years</i>	-0.0120 (.0022)	0.0010 <sup>1</sup> (.0060)		
<i>Race: Black</i>	0.1166 (.0357)	0.1553 (.0468)	0.1793 (.0351)	0.2092 (.0464)
<i>Race: Other Nonwhite</i>	-0.2619 (.0749)	-0.2397 <sup>1</sup> (.1276)	-0.1968 (.0790)	-0.1073 <sup>1</sup> (.1258)
<i>Disability</i>	0.8157 (.0310)	0.8113 (.0539)	0.8315 (.0308)	0.7733 (.0535)

<sup>1</sup>Not significant at 1 percent level.

for men and .091 for women. At any given age, mortality rates are higher for blacks and lower for other racial groups relative to whites, and disabled workers have much higher mortality rates.

Of most interest for present purposes are the income coefficients. Of the six income variables only zero income for males and cap income for females are insignificant. Mortality is significantly negatively related to log-income for men and women and to cap income for men and zero income for females. As a group, the three coefficients on average income, zero income, and number of years at the earnings cap (all measured over ages 46-60) are highly significant. The

chi-square values for the likelihood ratio tests when the variables are excluded from the regressions are 84.6 for men and 141.2 for women, with three degrees of freedom.<sup>9</sup>

The income effects are also large enough to change the implications about the role of race and birth year. Table 1 shows when the mortality equations are estimated with the income variables excluded, the coefficients on black race increase from 0.116 to 0.179 for males and from 0.155 to 0.209 for females. Apparently, much of the difference in life expectancy between the black and white racial groups can be attributed to income differences, a result which is consistent with Rogers (1992) and Menchik (1993). Also, with the income variables excluded, the birth-year coefficients rise in absolute value from 0.011 to 0.014 for males and from 0.005 to 0.010 for females, implying much of the upward trend in life expectancy is due to increases in real incomes.

Figures 1 and 2 highlight the estimated influence of income on mortality for males and females, respectively. The curves displayed in the figures are survivor rates, the predicted proportions of age-65 workers still alive at subsequent ages. The hypothetical individuals are white, non-disabled males and females born in 1920, compared at three different income levels. The middle-aged male wage worker is assumed to have earned an average of \$15,000 per year; females are assumed to have earned an average of \$8,000. Low- and high-incomes are about one standard deviation below and above the mean (\$5,000 and \$25,000 for males; \$1,000 and \$15,000 for females). We assigned the mean number of years at the cap (about 6 for males; a little over 1 for females) and assumed none of the workers had years of zero income.

Figure 1 shows the differences in income correspond roughly to one- or two-year differences in life expectancy for males. Reading horizontally from the 50 percent value, the

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<sup>9</sup>We tried a squared log(income) term to allow for more flexibility in functional form but it was insignificant in both the male and female equations.

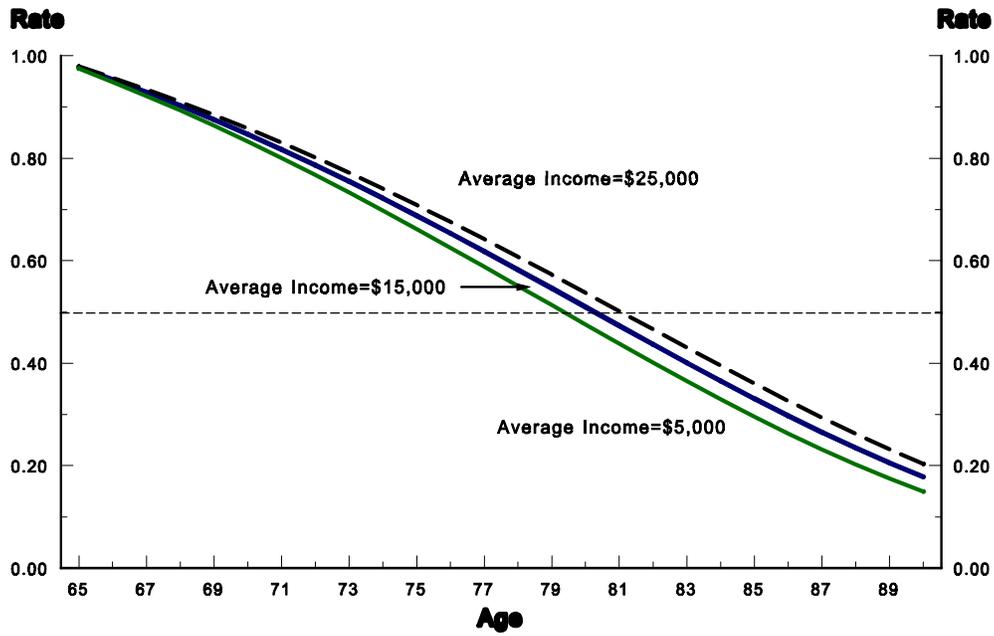
predicted median lifetimes are 79, 80, and 81 years for the low, middle, and high earners, respectively. For females, the income differences are slightly larger. Reading across the 50 percent value in Figure 2 shows predicted median lifetimes for females of approximately 83, 85 and 86 years for the low, middle, and higher earners, respectively. These differences in predicted mortality are certainly large enough that we would expect them to affect the estimated progressivity of Social Security benefits.

### **III. Mortality and the Progressivity of Social Security Benefits**

#### *IIIa. Social Security Benefits and Internal Rates of Return*

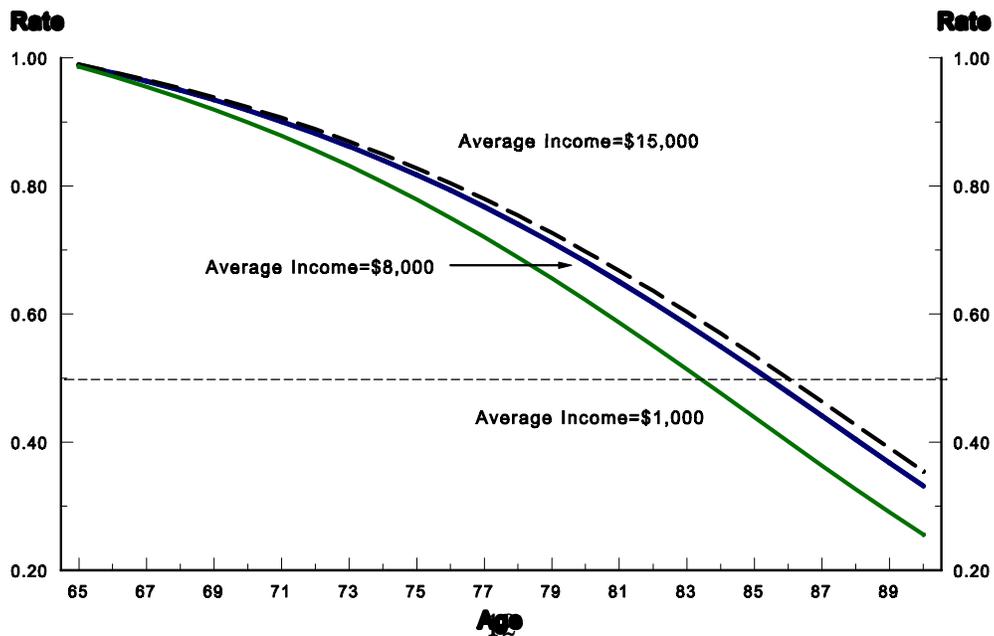
In this section we estimate lifetime Social Security benefits and rates of return for a subsample of our CWHS covered workers using mortality distributions from the parameter estimates shown in Table 1. For the purpose of estimating benefits and rates of return, the CWHS has considerable

Figure 1  
Estimated Survivor Rates for Males, 65+  
at Three Different Income Levels



advantages. It permits a more complete analysis of the Social Security contribution and benefit

Figure 2  
Estimated Survivor Rates for Females, 65+  
at Three Different Income Levels



base than has been possible in previous studies. No other available file contains longitudinal earnings data for such an extended period and such a large number of individuals. In contrast to earlier research, we are able to estimate the returns received by early cohorts based on actual program experience. In addition, our actual case histories include a wider range of situations than prior studies and cover the period through 1988, several years past the last major reform to the Social Security program.

For this paper, we have chosen to focus on the OASI program. We do not include disability benefits in program returns, and we exclude the DI portion of the Social Security (Federal Insurance Contributions Act, or FICA) contribution rate. However, we did not delete disabled individuals from our sample. Upon reaching age 65, disability beneficiaries are automatically recategorized as old-age beneficiaries, and their subsequent benefits are paid from the OASI trust fund. We include these OASI benefits as returns to their prior contributions.

Under the OASI program, those potentially eligible for benefits include retired old-age workers and certain of their dependents (spouses and/or children), and survivors of insured workers (widows and widowers, eligible children, dependent parents) who may or may not have been beneficiaries. In this paper, we analyze benefits paid to three principal beneficiary types: insured workers (old-age benefits), their spouses (spousal benefits), and surviving spouses (widow(er) benefits).<sup>10</sup>

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<sup>10</sup>Because the CWSHS provides birthdates only for insured workers and spouse beneficiaries, it would be highly problematic for us to forecast either the durations or levels of future benefit flows for other dependents. Therefore, we excluded those dependent benefits from our analyses, implicitly setting them to zero in all cases.

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Ignoring nonspouse dependent benefits will impart a clear downward bias to our rate of return calculations. However, the degree of bias is difficult to judge. Approximately four percent of OASI benefits are paid to, or on behalf of, survivors other than spouses and to children of current beneficiaries (Department of Health and Human Services (1991, Table 5A). Furthermore, many of the insured workers in this group, on whose account such benefits would be paid, died at relatively young ages, and so were below average in both contributions and benefits.

Insureds who were deceased in 1988 and who were no longer paying survivor benefits present especially difficult measurement problems. For these individuals, we do not observe any past or current benefit levels, nor can we identify the mix of beneficiaries, if any. We could not exclude them from our sample, however, without introducing serious biases. Deceased individuals, including those who died prior to entitlement, were therefore included in our computations if they satisfied our other editing criteria.<sup>11</sup>

The present value of lifetime Social Security benefits, SSB, from year of retirement, R, until year of death, D, is computed as the sum of expected benefits in each period, B<sub>i</sub>, discounted by the rate of interest, r to year T (1988 in our computations):<sup>12</sup>

$$(2) \quad \begin{aligned} \text{SSB} &= \sum_{i=R}^{T-R} [B_i / (1+r)^{T-i}] + \sum_{i=D}^T [B_i / (1+r)^{T-i}] \quad D \leq T \\ &= \sum_{i=R}^{T-R} [B_i / (1+r)^{T-i}] \quad D < T. \end{aligned}$$

Lifetime contributions, SSC, from the age of entry into covered employment, A, until retirement are computed as the sum of covered wages, W, times the (employer plus employee) Social Security contribution rate, c<sub>i</sub>, times the rate of interest to time T:

$$(3) \quad \text{SSC} = \sum_{i=A}^{R-A} [W_i c_i / (1+r)^{T-i}].$$

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<sup>11</sup>Further details of sample selection can be found in Duggan, Gillingham, and Greenlees (1993).

<sup>12</sup>If D < T the insured is deceased at the time of the CWSHS sample. Because the CWSHS contains actual benefit data only for the current year (1988 in our file), a crucial part of our work involved computing benefit histories using the information in the file and programming current and past benefit formulas, as detailed in the appendix.

The internal rate of return is that constant rate of interest  $r$  for which

$$(4) \quad \text{SSB} - \text{SSC} = 0.$$

### *IIIb. The 1917-22 Cohorts*

We focus our attention here on 8,344 workers (4,636 men and 3,708 women) born between the years 1917 and 1922, who received benefits under "new law," that is, under the specifications of the 1977 Social Security amendments. They include the group often referred to as the "notch" generation (birth years 1917-21), whose treatment relative to prior and subsequent cohorts was examined in Duggan, Gillingham, and Greenlees (1996). For present purposes, the 1917-1922 cohorts are a convenient group because they reached normal retirement age in the years just prior to 1988. Projections of future mortality rates, therefore, would be of critical importance to their estimated lifetime Social Security benefits. Also, the concentration on new-law beneficiaries implies our progressivity results should apply broadly to currently insured workers.

Several factors remain that limit the quantitative significance of our alternative assumptions about the relationship between income and life expectancy. First, even in the 1917-1922 subsample, a sizable percentage of covered workers had died by 1988. For these workers, benefits and rates of return are known values under any parameter assumptions. Second, our models apply only to insured workers and not to their spouses. We project the mortality rates of spouses, widows, and widowers using Census mortality data (see footnote 8).<sup>13</sup> For many married couples, this means our income-based life expectancies will not affect the point at which benefits

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<sup>13</sup>This applies not only to post-1988 mortality projections, but also to the imputation of death years for those spouses who were no longer living in 1988 (see the appendix).

cease, but will only determine the point at which the spouse stops receiving a spousal benefit and begins receiving a survivor benefit (and the total family benefit falls by approximately 33 percent).

### *IIIc. Simulated Lifetime Benefits and Rates of Return*

We simulated the benefit stream for each sample beneficiary by first drawing a random number from the uniform zero-one interval, then computing hazard rates and survivor probabilities for each year after 1988 using the parameters from Table 1. Death was assumed to occur in the year when the survivor probability fell below the random number. We then contrasted the sample present values of benefits, and real rates of return to prior contributions, in the two mortality models (income-adjusted and unadjusted) and at different levels of earnings.<sup>14</sup>

Our comparisons are presented in Table 2. In the first two columns are the estimated present values of returns, net of contributions, by income class. We defined four income groups based on income from age 46 to 60. The first consists of all individuals with zero incomes over that age range, separately classified because we expect total family income to be underestimated for this group. The next three classes are three equal quantiles of the sample, ranked by income (men and women are ranked and classified separately). The right-hand columns display real annual rates of return to contributions.

For males, the overall mean of net benefits is about \$52,000 under either mortality model. Both models indicate a 5.5 percent real rate of return for the male sample as a whole. Both also show a strong, although not monotonic, relationship between income and net benefits resulting from the combination of contribution and benefit rules. The male cohorts received very high real returns to their contributions, and by contributing larger amounts, higher-income workers received

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<sup>14</sup>We used the interest rates on the bonds held by the Social Security trust funds for computing present values. The Consumer Price Index CPI-U-X1 series, incorporating the current treatment of homeownership costs since 1967, was used for defining real returns.

**Table 2**  
**Social Security Net Benefits and Real Rates of Return**  
**Using Mortality Distributions Either Adjusted or Unadjusted by Income**

<i>Population</i>	<i>Income Class</i>	<b>Net Benefits (dollars)</b>		<b>Real Rate of Return (percent)</b>	
		<i>Income-unadjusted Mortality</i>	<i>Income-adjusted Mortality</i>	<i>Income-unadjusted Mortality</i>	<i>Income-adjusted Mortality</i>
<b>Men</b>	Zero	\$29,443	\$27,690	6.61	6.52
	Low	42,154	40,841	6.23	6.17
	Medium	62,798	62,528	5.59	5.58
	High	53,214	55,161	4.99	5.04
	All	51,927	51,985	5.47	5.47
<b>Women</b>	Zero	63,826	56,232	8.52	8.38
	Low	57,900	55,419	9.24	9.19
	Medium	55,891	56,727	7.66	7.70
	High	56,629	59,807	6.02	6.12
	All	57,380	57,229	7.36	7.36

greater benefits in present value terms. However, the income-adjusted mortality model gives somewhat smaller benefits than the unadjusted model for the first two (lower-income) classes, and larger benefits for the highest income groups. For the highest quantile, the estimated present value is \$55,161 in the income-adjusted model, compared to \$53,214 in the unadjusted model. In terms of the relative net benefits paid to the *A<sub>low</sub>* and *A<sub>high</sub>* income groups, the income-adjusted model predicts an average difference of over \$14,000, whereas the unadjusted model estimates only about an \$11,000 difference.

The direction of effect is similar in the rate of return columns. Again, the result of incorporating the effect of income on mortality is to slightly raise the estimated return to higher-income workers, and lower the return to lower-income workers. Both models show the real rate

of return falls as income rises, but the relationship is slightly weaker using the income-adjusted model. The unadjusted model results for males display a 1.24 percentage point difference in the rate of return between the low- and high-income groups, and the income-adjusted model estimates a 1.13 percentage point difference. (Comparing the zero- and high-income groups, the differences are 1.62 points for the unadjusted model and 1.48 points for the income-adjusted model.)

The results are quantitatively stronger for female workers. The unadjusted model shows flat or even declining net benefits for higher-income workers, though real rates of return decline for these workers. On the other hand, the income-adjusted model for women exhibits increasing net benefits for higher-income workers, resulting in a \$4,400 difference between the low- and high-income classes. The real rate of return is negatively correlated with income for women but less so in the income-adjusted model. The rate of return for lowest and highest income classes are significantly affected by the income adjustment to mortality.

### *IIId. The Marginal Effect of Income*

In Table 2, birth year, race, and other factors are not held constant in the group comparisons. Thus, the lower-income groups in the table may contain a greater proportion of disabled workers, and this is likely to affect the relative benefits earned by those groups. In order to control for these factors and to further highlight the extent to which the income-mortality relationship reduces benefit progressivity, we regressed individual benefits and rates of return against five dummy variables for the 1918 through 1922 birth cohorts, black and other race, disability, log-income, and a dummy variable for zero income.

Figures 3 and 4 show the predicted values from the income-adjusted and unadjusted male regressions at different levels of real income.<sup>15</sup> The marginal effect of income on net benefits in

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<sup>15</sup>The regressions use the mean sample values of dummy variables. They also exclude sample workers who died before receiving benefits since their rates of return are undefined. Also, these workers died at ages below

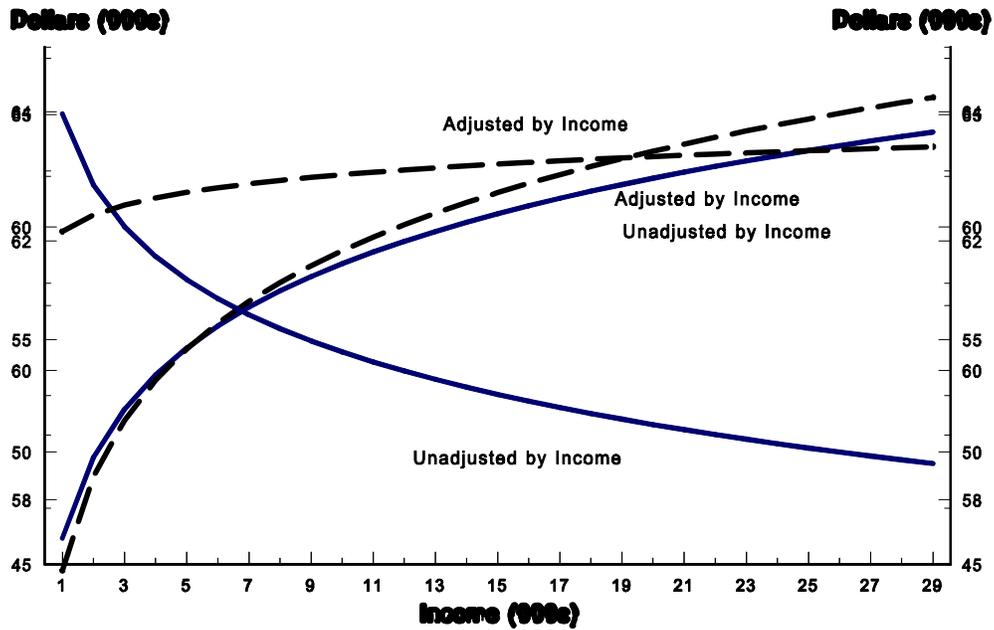
Figure 3 is smoothly positive in both models, but stronger in the income-adjusted model. The marginal effect of income on rate of return in Figure 4 is smoothly negative in both models, but weaker in the income-adjusted model.

Figures 5 and 6 display comparable data from the female regressions. Notice the declining net benefits in Figure 5 for the unadjusted model, which corresponds to the findings we reported in

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those for which our mortality projections come into play.

**Figure 5**  
**Net Present Values of Social Security Benefits for Males**  
**With Mortality Either Adjusted or Unadjusted by Income**



**Figure 6**  
**Real Rates of Return to Social Security Contributions for Males**  
**With Mortality Either Adjusted or Unadjusted by Income**

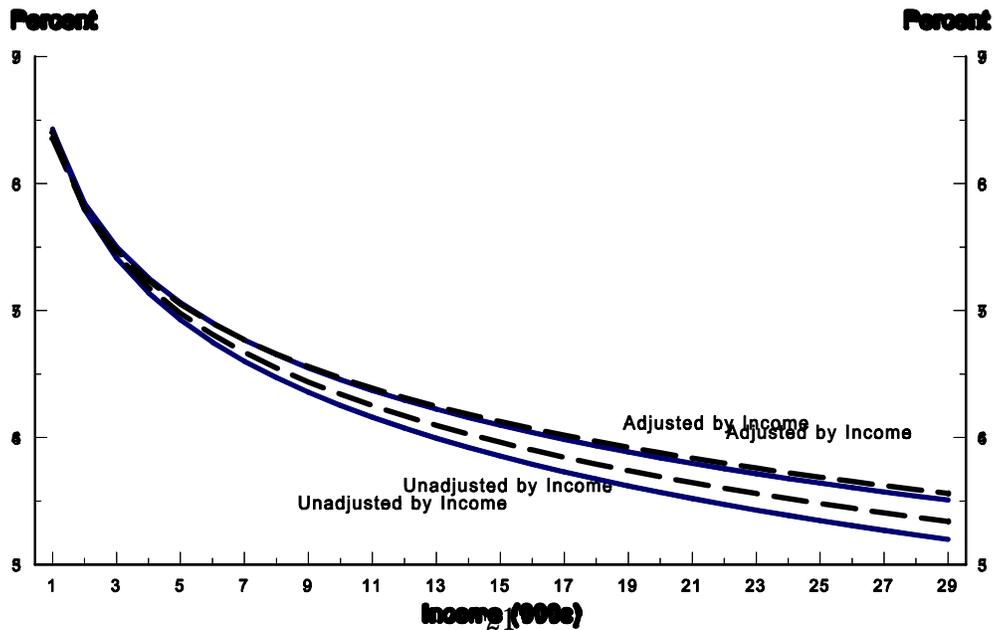


Table 2. The income-adjusted model reverses this result, showing a smoothly positive marginal effect of income on net benefits, as in the male model. Figure 6 reveals a uniformly negative effect of income on the rate of return for females, but a weaker one in the income-adjusted model.

Thus, for both genders and both for net benefits and real rates of return, the income-adjusted mortality model gives noticeably less progressive results. However, the strongly progressive relationship between income and the return to Social Security contributions is qualitatively unchanged.

#### **IV. Summary and Conclusions**

In this paper we addressed two fundamental questions. First, do higher income persons live longer? To answer this question, we employed a sample of over 44,000 insured workers born between 1895 and 1923 to examine the impact of income on life expectancy. In this effort we employed a logit model of the probability of death in a given person-year over age 65. The results provided strong evidence of a causal effect of income, even when race, sex, disability, and birth year were held constant.

Second, does recognition of the income-mortality relationship affect the progressivity of Social Security benefits? This was the question posed in the title of the paper and provided the motivation for the first question. To learn the answer we simulated the future mortality experience of a subsample of 8,344 CWHS workers born during the period 1917 to 1922, and compared their simulated benefits obtained using logit models of mortality either adjusted or unadjusted by income. We found the income-adjusted model produced less progressive results, in the sense that it yielded relatively greater net benefits and real rates of return for higher-income workers. Thus, Friedman was right in his 1972 statement about income, life expectancy, and Social Security net returns. Empirically, the effect is small, however. Our results show the effect of income-adjusted

mortality on returns to Social Security contributions does not reverse the conclusion that Social Security net returns are strongly progressive.

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