SIEPR Discussion Paper No. 14-009

Changing Mortality Rates and Income Inequality among the U.S. Elderly

By

Gary Burtless

Changing Mortality Rates and Income Inequality among the U.S. Elderly

by

Gary Burtless[1]

THE BROOKINGS INSTITUTION

October 6, 2014

TWO TRENDS HAVE INCREASED disparities in well-being among aged and near-aged Americans. First, money income inequality has increased over time. Among members of families headed by a person age 62 or older, inequality as measured by the Gini coefficient has increased 11 percent since 1979. This increase has been driven by some of the same factors that have pushed up inequality among the nonelderly. These include growing gaps in labor incomes and increased concentration of capital income. Second, life spans have become more unequal. A venerable literature has established that there are significant differences in life expectancy between people with high and low socioeconomic status as measured by such indicators as income and educational attainment. New studies also suggest that the differential has widened in the United States over the past three decades, reversing an earlier trend toward greater equality. A few studies even suggest that much of the recent increase in expected life spans is concentrated among those with above-average incomes, and that life expectancy may be roughly constant or even declining for Americans with lower status.

Trends in differential mortality uncovered in recent research raise profound questions about the equity of old-age pension formulas. The Social Security retirement-worker pension provides a basic benefit at the normal retirement age, known as the Primary Insurance Amount or PIA. The formula for this pension is highly redistributive. It provides a more generous replacement rate for low-lifetime-wage workers than for workers with high average earnings. This kind of redistribution may be necessary to compensate low-wage workers for their shorter expected life spans. Retired workers' actual Social Security benefits are determined by (a) their PIA, and (b) the actuarial factors used to adjust the monthly pension to reflect early or late benefit claiming. Workers who claim benefits at the earliest entitlement age, 62, receive reduced benefits; workers who delay claiming benefits until the latest claiming age, 70, receive monthly

[1] I gratefully acknowledge the crucial research contributions of my colleagues Barry Bosworth, Mattan Alalouf, and Kan Zhang, all of whom played critical roles in creating and analyzing the research files. I also gratefully acknowledge the generous research support provided by the Sloan Foundation.

payments that are considerably higher than their PIA. Coincident with a trend toward later retirement, after 1990 we have also seen a trend toward later benefit claiming for Social Security retired-worker benefits (Bosworth and Burtless 2010). The workers who delay benefit claiming receive bigger monthly pensions as a result. If these workers had better-than-average earnings during their careers, the delay in benefit claiming increased the gap between their monthly Social Security benefits and the monthly benefits received by lower wage workers.

Differences in mortality mean that, for any given age at which benefits are claimed, high-wage workers can expect to collect benefits longer than low-wage workers who claim benefits at the same age. If gains in expected life spans are increasingly concentrated among high-wage workers, we may not want to ask less affluent workers to bear a large share of the financial burden of an aging society. A common suggestion to deal with funding shortfalls in Social Security and Medicare is to lift the age of eligibility for benefits. This policy would make sense if the gain in expected life spans is enjoyed equally by rich and poor alike. It seems less equitable to ask low-wage workers to wait longer for retirement benefits when a disproportionate share of gains in life expectancy have been enjoyed by the affluent. In view of changing relationship between workers' average lifetime earnings and their chances of surviving into late old age, how can we recalibrate the PIA formula and the actuarial adjustment for delayed benefit claiming to protect the interests of low-wage workers? This essay tries to expand the base of knowledge with which we can answer this questions.

The remainder of the essay is divided into two parts. The first section examines trends in old-age income inequality in the past three decades. In particular, I use evidence on money income obtained in the Census Bureau's annual CPS income survey to examine inequality within narrow age groups in the population. I show how inequality within these narrow subpopulations has changed over time and how inequality within given birth cohorts evolves as the birth cohort grows older. Money income inequality has risen over time. This is the case for the U.S. population generally and also within narrow age groups in the middle-aged and aged populations. However, the growth of inequality has differed in the aged and nonaged populations. First, inequality has increased faster among the nonaged than among the aged. Second, at least in the lower half of the income distribution inequality begins to decline starting at age 62 when workers and their dependent spouses become eligible for early retired-worker benefits.

The second main section presents new results on the growing mortality differential between Americans based on their average Social-Security-covered earnings. With my Brookings colleagues Barry Bosworth and Kan Zhang, I have organized a large data file containing information from the Census Bureau's Survey of Income and Program Participation (SIPP) matched to Social Security lifetime earnings records and Social Security mortality records. These data permit us to analyze determinants of mortality within a large sample SIPP respondents born between 1910 and 1956 over the period from 1984 through 2011. Besides obtaining access to these confidential data, the most challenging part of the research is devising measures of socioeconomic status that permit us to make evenhanded comparisons between generations born over a 46-year time span. This is a challenge because the measures of annual earnings contained in the Social Security Administration (SSA) files were subject to different reporting limits during the 10-year age spans we use to estimate average earnings for successive birth cohorts. In this initial analysis of the data, I offer results based on one way of dealing with the limitations of the data.

Income inequality across age groups

Money income inequality has increased noticeably since 1979 (DeNavas-Walt and Proctor 2014, Table A-2). Although the amount of increase in inequality differs depending on the measure of income used, there is little question that inequality has risen under virtually all measures of both pre-tax and after-tax income (U.S. Congressional Budget Office 2013). Theories to explain the increase abound. One factor pushing up inequality is the rise in wage disparities. This factor primarily affects working-age breadwinners and their dependents, because labor income constitutes an overwhelming share of these families' incomes. Rising wage inequality can also boost inequality in old age to the extent that labor income inequality leads to increased inequality in family savings and breadwinners' pension accumulations. Aged Americans are, however, far more dependent on government transfers, including Social Security benefits, than are the nonaged. Since transfers tend to represent a larger percentage of the incomes of families with low incomes, the fact that public benefits have been largely protected over the past three decades means that incomes of the low-income elderly have fared better than those of the low-income working-age population.

In order to trace inequality trends in the aged and nonaged populations, I use money income data collected in the Census Bureau's Annual Social and Economic Supplement, usually referred to as the March CPS. The tabulations presented here cover every third calendar year from 1979 through 2012. The income measure that I use is based on the standard Census Bureau definition of "money income." It is derived from respondents' reports of pre-tax income from wages, self-employment, capital income sources, and cash government transfers, including Social Security and public assistance. It excludes in-kind benefits such as housing assistance, food stamps, and government- and employer-provided health insurance. The public use version of the CPS file uses an inconsistent method for top-coding high income amounts reported by respondents. In effect, the top-coding procedure truncates reported incomes much more severely in the 1980s and early 1990s compared with later years. To circumvent this problem I replaced the original Census Bureau top codes with alternative codes proposed by analysts with access to the uncensored data (Larrimore et al. 2008).

In order to divide the population into age groups, I classified each family by the age of the head of family or, in the case of married-couple families, the older of the head and the spouse of the head. Single-person households and unrelated individuals are also classified by the person's age, and they are designated as family units. If more than one family resides in the same household, each family is separately classified by the age of its head. I ranked families according to their family-size-adjusted incomes and then used these family ranks to determine the income ranks of people who were members of the families. These person ranks are based on their families' rank in the size-adjusted income distribution.[2] Inequality is ascertained by calculating standard measures of income disparity for persons rather than families based on each person's family size-adjusted income.

The trend in size-adjusted income inequality for the entire noninstitutionalized population is shown in Figure 1. The Gini coefficient of inequality increased from 0.379 to 0.482, or 27 percent, between 1979 and 2012. Almost nine-tenths of the increase occurred between 1979 and

[2] The family size adjustment is intended to determine families' income rank by their "equivalent" or "size-adjusted" incomes, that is, their family income adjusted to reflect the effects of family size. The adjustment used is to divide each family's unadjusted income by the square root of the number of family members. This adjustment implies that a family consisting of four members requires twice as much income to have the same "equivalent" income as a household containing just one member. Note that each of the income quantiles contains an equal number of *persons* rather than an equal number of *families.*

2000. Inequality as measured in the March CPS has increased much more slowly since 2000. Figure 2 shows separately the trends in inequality among individuals in families headed by aged and nonaged adults. I define an "aged" family head as someone who is at least 62 years old. Sixty-two is the earliest age at which Americans can claim a retired-worker benefit under Social Security. In 1979 inequality in aged families was considerably higher than among people who were members of nonaged families. Since then, however, inequality has increased much more steeply among people in nonaged families. Between 1979 and 2012 inequality in this population increased from 0.368 to 0.486, or an astonishing 32 percent. Size-adjusted income inequality increased just 11 percent among people who were members of aged families. Inequality in aged families is now modestly lower than it is in families with a nonaged head.

The classification of families by the age of their heads permits us to measure inequality within even narrower age groups. I divided families into three-year age groups starting at age 47 and ending at age 79. (Families headed by a person 80 or older are placed in a single age group, because the number of such families is small.) Figure 3 shows Gini coefficients within these narrower age groups over the period from 1979 to 2012. To make the results clearer, I have averaged results for two groups of calendar years: 1979, 1982 and 1985 at the start of the analysis period and 2006, 2009, and 2012 at the end. In addition, the chart shows the average value of the Gini coefficient over all 12 of the years in the analysis sample, 1979-2012. The tabulations show that inequality typically rises from age 47 through 64, but then either declines or remains roughly unchanged. The results are cross-sectional, that is, they show the age profile of money income inequality within single calendar years. They suggest that the cross-sectional pattern of inequality has changed over time. Whereas income inequality peaked among persons in families headed by 74-to-76 year-old family heads in the late 1970s and early 1980s, in more recent years the peak level of inequality has been attained by families headed by someone who is between 62 and 67 years old. Income inequality is nowadays lower among people in families headed by a person 77 or older than it is among people in families headed by someone who is younger. Inequality has increased in the oldest families, but it has increased far less than it has in younger families.

Figure 4 highlights this pattern by showing the percent change in the Gini coefficient among families classified by the age of the family head. Whereas families headed by someone in their late 40s or early 50s saw inequality rise by about one-quarter, families headed by someone past 67 saw a much more modest increase. It is plausible to think the change in the age pattern of inequality is linked to the importance of labor income in families' total income. As labor income inequality has increased, families that largely depend on labor earnings for income have experienced rising inequality. At ages past 58, and especially past 67, labor income is gradually replaced by retirement income sources, such as Social Security, Supplemental Security Income, and pensions. Because of the redistributive tilt in the benefit formula, Social Security benefits are much more equally distributed among families that receive them than are labor incomes. Nonetheless, families headed by someone between 62 and 74 are increasingly affected by trends in wages. Since the early 1990s U.S. workers have been delaying their retirements and increasing the share of their incomes derived from labor earnings (Bosworth and Burke 2012). Considerable evidence suggests that retirement delays have been especially common among well-educated, highly compensated workers (Burtless 2013). Labor income is much less common among families headed by a person past 74.

Because we have classified families into 3-year age groups by the age of the family head and we have tabulated inequality statistics every third year, it is possible for us to trace out the trend in inequality for individual birth cohorts. Figures 5 and 6 show the results of these tabulations. Instead of measuring inequality with the Gini coefficient, the two charts show the separate trends in inequality in the bottom half and in the top half of the size-adjusted income distribution. Figure 5 shows the ratio of size-adjusted incomes in the 50th and the 10th percentiles of the income distribution; Figure 6 shows the 90/50 income ratio. The two charts show trends in inequality in single birth cohorts as the cohort ages. The tabulations displayed are for three birth cohorts: family heads born between 1915 and 1917, heads born between 1930 and 1932, and heads born between 1945 and 1947. The cohort born in 1930-1932 was 47-49 years old in 1979, and it was 77-79 years old in 2009. Therefore, the entire trajectory of inequality from age 47-49 through 77-79 can be traced out for this cohort. The younger and older cohorts are observed for a smaller number of ages between 47 and 79.

Figures 5 and 6 show, not surprisingly, that inequality has been higher at given ages for the younger cohorts compared with the older ones. The differences are greater, however, for the 90/50 income ratio than for the 50/10 ratio. A reasonable inference is that top end inequality has contributed more to the increase in overall inequality than has an increase in bottom end inequality. The two figures also show a contrasting pattern of inequality change after age 61.

Whereas the 90/50 income ratio continues to increase after 61, the 50/10 ratio declines noticeably. This divergence is probably explained by the importance of government transfer benefits in holding up and even raising the incomes of older Americans once they attain age 62. In the bottom half of the old-age income distribution labor income is comparatively unimportant, and therefore the trend in contemporaneous wage inequality plays a limited role in explaining the gap in income between a family in the exact middle of the distribution and a family at the 10th income percentile. Neither family has much labor income, especially after age 65. In contrast, labor income remains important for families in the top of the old-age income distribution. Moreover, as noted above, labor income has become increasingly important for top ranking families as breadwinners have delayed their exit from the workforce.

In sum, the trend toward greater inequality in the working-age and general population can also be seen in families headed by an aged person. However, the increase in inequality has been smaller in the aged population, especially in the population past 74. One reason is that labor income is less important, and government transfer income more important, for these families than for the young. Delayed retirement, particularly among better educated and more highly compensated workers, has boosted the importance of earned income among the "young old," and we can see evidence for this in a growing gap between the incomes of middle-income and top-income aged families headed by someone between 62 and 74.

The rising mortality gap between the affluent and poor

Money income inequality is not the only indicator of disparities in well-being. Another indicator of special significance to the elderly is the inequality of expected life spans. Evidence of a widening in the difference in mortality by social and economic status has been found in a number of recent studies. Meara, Richards and Cutler (2008) and Olshansky and others (2012) analyze death certificate data, using educational attainment as a measure of status, and find a sharp rise in inequality. Waldron (2007) uses administrative records containing information on career earnings and age at death to establish a similar pattern for men covered by Social Security. Singh and Siahpush (2006) provide additional evidence on changes in differential mortality using county-level information from the decennial censuses. They construct county-level indexes of SES linked to death records by location. In a new analysis, Bosworth, Burtless, and Zhang (2014) use longitudinal data from the Health and Retirement Study (HRS) combined with Social

Security earnings and death records to estimate the relationship between mortality and lifetime earnings and other indicators of social and economic status.

The SIPP sample. In this section I extend the earlier analyses by estimating trends in mortality differences using SIPP longitudinal survey files that have been matched to earnings and mortality records in the Social Security administrative files. Our dataset contains records for SIPP respondents born between 1910 and 1956 who were sampled in the 1984, 1993, 1996, 2001, and 2004 panels. We were able to successfully match about 80 percent of the respondents to their corresponding Social Security earnings and death records, yielding a total sample of over 70,000 men and over 81,000 women.[3] Among these, about 100,000 respondents were "married, with spouse present" at the time of the SIPP interview. We were able to successfully match slightly less than 98 percent of these married respondents to their spouse's Social Security record. We then created a person-year dataset, in which each respondent enters the sample in the year corresponding to their initial SIPP interview (beginning in 1984) and remains in the sample until the year of their death or until 2012 (the last year for which we have reliable death data).

In estimating mortality functions for men, we used as an indicator of economic status each male's own nonzero earnings between the ages of 41 and 50. Earnings records for those ages are available for almost 55,000 male respondents, providing us with 680,000 person-year observations of potential mortality for men. About 15,000 of the 55,000 male respondents died over the observation period.

It is less obvious how to measure the social and economic status of married women. A woman's own earnings can provide a clear indication of social and economic status for never-married women, but may offer a poor indicator of status for married women, especially those who earned relatively little during years they were rearing children. My indicators of status for women are therefore more complicated than the earnings indicators for men. For never-married women I use average nonzero earnings of the women when they were between 41 and 50 years old. Because own earnings may be a poor indicator of economic status for married women, I preferred to use an indicator that combined the earnings of the woman and her spouse. In

[3] We assume in the remainder of the paper that the 20 percent of SIPP respondents who could not be matched to their Social Security records are randomly selected. Whether or not this is true is an open question, but not one that will be examined here.

specifications that include a combined spouse earnings measure, I was therefore forced to exclude all married or formerly married women who could not be matched to a spouse's earnings record. This exclusion affects all the women who entered the SIPP sample as widowed, divorced or separated respondents. This exclusion reduces the sample of women to 37,000 unique respondents, or just over 500,000 person-year observations. Among these respondents, about 6,600 died during the observation period.

Another commonly used indicator of social and economic status is educational attainment. Information on schooling is available for all respondents who are 15 or over at the time of the SIPP survey. When we used this indicator of status we were able to use larger samples to examine the impact of status on mortality. This larger estimation sample contains a total of 710,000 person-year observations for men and over 850,000 person-year observations for women. In this sample there were a total 16,000 deaths among men and more than 15,000 deaths among women. Details about the sample of matched SIPP and Social Security records are reported in Table 1.

Measuring workers' earnings ranks in a consistent way. A crucial problem with the earnings records maintained by SSA is that earnings reported in the file are often capped at a maximum annual amount that is equal to the maximum earnings subject to the Social Security payroll tax in the year.[4] The problem is especially severe for male earnings reported in the 1950s and 1960s, when the taxable earnings cap was low in relation to the annual wages earned by prime-age men. Because the SIPP sample was not selected until 1984 or a later year, it is impossible to deduce from the earnings records in the matched SIPP-SS file the exact importance of the earnings cap on successive birth cohorts. That is because the respondents enrolled in the SIPP sample do not represent a random cross section of the men in a particular birth cohort who earned Social-Security-covered earnings. Men who failed to survive to 1984 are of course excluded from the SIPP sample.

To address this problem we turned to the SSA Earnings Public-Use File (EPUF). That file contains summary earnings information on individual-level Social-Security-covered earnings

[4] After 1980 the administrative records contain information about earnings above the taxable earnings cap. In earlier years, however, a worker's earnings above the cap must be imperfectly inferred using information about the quarterly pattern of earnings reported by a worker.

before 1951 as well as individual-year data on covered earnings between 1951 and 2006 (Compson 2011). The data set includes information for a random sample of 3.13 million Americans who had covered earnings in at least one year between 1951 and 2006. The data in this file can be used to determine the exact percentile of the male earnings distribution that is just below the taxable earnings cap for each birth cohort when it is between 41 and 50 years old. We determined that the lowest earnings cap was slightly above the 31st percentile of earnings for the male birth-year cohort that experienced the lowest annual earnings cap (relative to the cohort's annual earnings distribution). In order to construct an indicator of workers' positions in the male earnings distribution, we therefore used workers' earnings up to the 31st percentile in the EPUF file. The 31st percentile earnings amount was estimated using the EPUF earnings for a given birth cohort and separately for each year of age for that cohort between 41 and 50. These estimates of a low earnings level are based on reports in the EPUF file for all males in the population who reported nonzero Social-Security-covered earnings in the year. They are therefore unaffected by the subsequent mortality experience of men in the cohort.

After identifying the 31st percentile earnings amount for each male birth cohort and each year of age between 41 and 50, we constructed alternative classification schemes to identify "persistent low earners" or "likely low earners." Workers in the matched SIPP-SS sample who had nonzero earnings equal to or less than the 31st percentile amount in all years for which they reported nonzero earnings between ages 41 and 50 can be reliably classified as "persistent low earners." Slightly more than one-quarter of the men in our SIPP sample met the earnings criteria to fall in this category. Our more expansive definition of a "likely low earner" requires that the worker earn less than the 31st-percentile amount in at least half of the years between ages 41 and 50 in which he reports nonzero earnings. This class of earners constitutes 43 percent of the male workers in our sample. Clearly it would be preferable to classify workers by their average earnings over a longer part of their careers and using all earnings, rather than just Social-Security-taxed earnings. However, the limits of the Social Security earnings file make this impossible. The classification scheme used in this paper permits us to identify low earners in a consistent way given the varying limits on the annual income amounts reported in the Social Security earnings file.

Women's reported earnings are much less affected by the maximum taxable earnings amount. In years when the taxable cap was low relative to economy-wide average wages – 1951 through the mid-1970s – women's earnings were also comparatively low. As women's average earnings increased, so too did the annual earnings cap. As a result, women's annual wages up through the 80th earnings percentile are observed, even in the calendar year with the lowest earnings cap relative to the female earnings distribution. Thus, it is easier for us to classify women earners by their exact position in the female earnings distribution. We simply counted women's annual earnings up through the 80th percentile and then calculated the average nonzero earnings amount between ages 41 and 50. Within each birth cohort we then ranked women by their position in the distribution of nonzero average earnings. In specifications that used women's own earnings to indicate their social and economic status, we used straightforward indicators of their earnings rank within their birth year cohort. For example, a "low earner" might be one whose average nonzero earnings placed her in the bottom half of her cohort's earnings distribution.

It is more complicated to construct a consistent indicator of low household earnings. The fact that reported average male earnings in the 1950s and 1960s have a much lower cap compared with earnings reported in 1977-2012 means that we cannot simply add the earnings of the two spouses to construct a valid and consistent indicator of their combined earnings. Our solution is to combine our separate indicators of low earnings status for the two spouses to create a composite indicator of low household earnings. For example, married women with a spouse who earns less than the 31st percentile earnings amount are always classified as having low household earnings, as are all unmarried women whose earnings are less than the 67th percentile of average nonzero earnings among women born in the same year.

Measuring educational attainment as a consistent indicator of status. Because of the difficulty of obtaining reliable measures of past income and earnings, it is common to use education as an alternative indicator of social and economic status. This variable is unquestionably linked to both income and status, and it is easily ascertained. For determining the relationship between status and mortality experience within a single birth cohort, educational attainment has powerful practical advantages. However, it offers a less reliable way to measure the changing influence of status on mortality experience across successive generations. Educational attainment has risen considerably over time. In the 1962 Current Population Survey, 58 percent of the men who were between 48 and 52 years old reported they had not completed high school; just 9 percent reported they had completed college. In the 2010 Current Population

Survey, only 11 percent of 48-52 year-old men reported they had failed to complete high school; 29 percent reported they had obtained a college degree. Clearly, the lack of a high school diploma was an indicator of much more serious disadvantage for 48-52 year-old men in 2010 than it was in 1962. Completion of college was a more marked indicator of social and economic advantage in 1962 than it was in 2010. If we find that failure to complete high school is associated with a much bigger increase in mortality in 2010 compared with 1962, we should hardly be surprised. Adults who failed to complete high school represented a much smaller and economically more disadvantaged population in 2010 compared with 1962.

Rather than use educational attainment as a direct measure of a person's status, I used the person's relative educational attainment within the person's birth cohort. To calculate a person's educational rank, I used annual Current Population Survey files from 1962 through 2013 to estimate the exact distribution of reported educational attainment in the population between 48 and 52 years old. I assume the reported distribution of schooling for this population approximates the schooling distribution of the population that is exactly 50 years old in the calendar year of the CPS interviews. In estimates shown below I report the results when matched SIPP-Social Security respondents are divided into below-median and above-median educational attainment groups. This classification depends on the reported distribution of schooling attainment for successive birth cohorts as reported in the CPS interviews. A problem with this classification scheme is the large number of persons in the CPS with tied educational ranks. In 1980, for example, 34 percent of men aged 48-52 had not completed high school, and 31 percent had obtained a high school diploma but no education after high school. This means that 31 percent of men had schooling attainment equal to exactly the median attainment level. In a classification scheme that divides men into below-median and above-median groups it is impossible to place all of the high school graduates in one group or the other. An ideal solution to this problem would be to further classify respondents by their class rank in high school, assigning those with lower class rank to the lower educational attainment group and those with higher rank to the higher attainment group. Since class rank data are not recorded in either the CPS or SIPP, I used the less satisfactory method of assigning a random fraction of high school graduates to the lower group, with the percentage determined by the share needed (according to the CPS survey) to place half of the cohort in the below-median attainment group.

Results. I estimated a parsimonious discrete-time logistic model to summarize observed mortality patterns in the matched SIPP-Social Security sample:

$$(1)\ \log (h_{it} / (1 - h_{it})) = \alpha_0 + \beta_1\, Age_{it} + \beta_2\, (Birth\ Year_i - 1900) + \beta_3\, Low\ Status_i + \beta_4\, (Birth\ Year_i - 1900) \times Low\ Status_i + \beta_5\, FYr_{it}\ ,$$

where $h_{it} = Pr(Y_{it} = 1 / Y_{it-1} = 0)$ is the hazard that person *i* will die in year *t*;

Low Status $_i$ = 1 if person *i* is in low status (schooling or earnings) group;
= 0 otherwise; and

FYr_{it} = 1 if year *t* is the first year person *i* is included in the HRS sample
= 0 otherwise.

(FYr_{it} is included in the specification to reflect the fact that respondents are exposed to the risk of dying for less than a full 12 months in the first calendar year of their enrollment.) The impact of low status on a worker's mortality is captured by the coefficients β_3 and β_4. If the impact of *Low Status* increases or shrinks in successive birth cohorts, the size of this impact will be reflected by β_4. If the mortality differential due to low economic and social status is growing, β_4 will be positive.

Results based on male earnings. Equation 1 was initially estimated separately for all men in the SIPP-Social Security sample described above, namely, males with positive nonzero earnings between ages 41 and 50 who were born between 1910 and 1956 and who were aged between 49 and 91 in the years of observation included in the estimation. The measure of *Low Status* used is based on the definition of a "likely low earner" described above. This definition requires that workers earn less than the 31st-percentile amount in at least half of the years they have earnings between ages 41 and 50. Recall that this class of earners represents 43 percent of the male workers in our sample.

Rather than present parameter estimates, I will present the estimated findings in charts that show the mortality rates or mortality-rate differentials associated with *Low Status*. Figure 7 shows the implied mortality rates predicted by the parameter estimates for the entire male sample. The top panel, Figure 7a, shows predicted age-specific mortality rates for low- and high-earning men born in 1925. The horizontal axis shows the men's age, and the vertical axis indicates the probability a surviving male will die at the indicated age. Men classified as low earners clearly face elevated mortality risk at every year of age past 50. The lower panel, Figure 7b, shows the implied mortality probabilities for low- and high-earning men born 20 years later, in 1945. At each year of age the risk of death has declined, but it has declined proportionately faster for high-earning compared with low-earning men. The estimated jump in the mortality differential is highly significant. (In this large sample, the *p*-value of the coefficient on β_4 is 0.0001.)

The results presented in Figure 7 show mortality predictions for every year of age between 50 and 90. However, for neither birth cohort did we directly observe mortality rates over that entire span of ages. Figure 8 shows the predicted mortality rates for high-earning men born in four different years, 1925, 1930, 1935, and 1940. In this case the predictions displayed are correspond with the ages for these cohorts where mortality is actually observed in the SIPP-Social Security sample. There are a range of ages, from 59 through 70, in which we directly observe mortality rates for all four cohorts. As expected, age-specific mortality rates decline in successive cohorts. A chart showing mortality rates in the same cohorts but for low-earning males would show higher mortality rates and a lower rate of decline across successive cohorts.

Figure 9 shows the implied ratio of the mortality rate in the low-earning group compared with high-earning males across successive birth cohorts. In the cohort born in 1920, the ratio of the age-specific mortality rate among low-earning males to that among high-earning males is 1.19, implying that at a given age the low-earning male is 19 percent more likely to die than a high-earning male. In the 1935 birth cohort the mortality rate ratio has risen to 1.60, and in the 1950 cohort it is 2.17. The fact that that the age-specific mortality rate is constrained to increase by a fixed proportional amount at every age and along a fixed linear path across successive cohorts represents a weakness of the specification.

The large sample size permits re-estimation of Equation 1 within narrower age groups to pinpoint the age ranges where widening mortality differentials are largest and most significant. I performed the re-estimation within overlapping 7-year age groups (49-55, 52-58, 55-61, 58-64, 61-67, etc.). Figure 10 shows predicted age-specific mortality rates for high-earning men born in 1925, 1930, 1935, and 1940. Each highlighted dot along the lines shows the central point estimate of estimated mortality within the age group, with the mortality rate calculated at the mean age in the subsample. The estimates correspond exactly to the more constrained estimates for the entire sample displayed in Figure 8. Note that the two figures present exactly the same picture of mortality improvement across successive birth cohorts.

Figure 11 shows mortality rate differentials between low and high earners in successive cohorts and at different ages. The differential is measured as the ratio of the age-specific death rate in the low-earning male group compared with the high-earning group. Again, each highlighted dot along the lines shows the central point estimate of this ratio in a birth cohort and within an age group, with the mortality rate ratio calculated at the mean age in the subsample. Even in this more flexible specification it is plain that the mortality differential has widened over time. The parameter that captures this widening is the coefficient β_4, and this coefficient is highly statistically significant for every age group between ages 49-55 and 76-82. The fact that the differential has not widened significantly at ages past 82 may reflect our small sample sizes after that age or the fact that low-earning men who survive past 82 are highly selected from a very healthy population.

Results for women based on combined spouse earnings. As noted above, it is more difficult to use the Social Security earnings records to construct a reliable indicator of married women's social and economic status. Some married women who have a low rank in the female earnings distribution may nonetheless have a high rank if their combined family earnings, rather than their own earnings, were used to ascertain their social and economic status. In results I shall not discuss further, I find that classifying women by their own earnings yields results that are similar to those I obtained for men. Women in the bottom half of the female earnings distribution (within their own birth cohorts) have higher mortality rates than women in the top half of the distribution, and the mortality rate differential has widened over time. The estimated increase in the differential is highly statistically significant if all women earners are included in the estimation, but when the sample is restricted to narrower age groups the effect is significant mainly for women between the ages of 57 and 80. At younger and older ages the effect is smaller or the samples are too small to yield a precise estimate of β_4.

The more interesting results are obtained when we assess a married woman's social and economic status using an indicator of the combined position of the woman and her spouse in the female and male earnings distributions, respectively. In obtaining the results reported below, I classified never-married women has having low family earnings if their own average nonzero earnings between 41 and 40 was less than the 67th percentile of earnings for women born in the same year. For married women, regardless of their own earnings, who had a spouse with earnings above the 31st earnings percentile for men born in the same year in all years between ages 41 and 50, the woman was classified as a member of a high- or average-earnings family. For other married women, the classification of the woman's status depended on the percentile rank or earnings pattern of the two spouses. In all 50 percent of never-married women were classified as low earners as were 47 percent of married women in the sample. (Ever-married women were dropped from the estimation sample if they could not be matched to a spouse's Social Security earnings record.)

Equation 1 was estimated using this indicator of low earning status for the sample of married and unmarried women who were born between 1910 and 1956, who were aged between 49 and 91 in the years of observation included in the estimation, and who had the required earnings data needed to calculate their low earnings status. Figure 12 shows the excess mortality rates for low-family-earnings women compared with average or above-average women in four birth cohorts. Note that the estimates displayed do not indicate the age-specific mortality rates of women in the low-family-earnings group. They show the *increase* in age-specific mortality rates relative to higher earning women born in the same year and at the same age. The results indicate that the mortality rate differential has increased over time. At each year of age, the mortality rate difference is bigger more recent birth cohorts compared with cohorts born in earlier decades.

The results displayed in Figure 12 reflect predictions made when women of all ages past 48 and younger than 92 are included in the estimation sample. As noted above, the parsimonious specification of equation 1 could needlessly constrain the estimates of the widening mortality differential. We can lessen the impact of this constraint by re-estimating equation 1 within samples restricted to a narrower age range of observations. I performed the re-estimation within overlapping 9-year age groups (49-57, 53-61, 57-65, 61-69, etc.). Figure 13 shows predicted age-specific mortality rates for women in the average- and above-average earning group born in 1925, 1930, 1935, and 1940. Each highlighted dot along the lines shows the central point estimate of estimated mortality within the age group, with the mortality rate calculated at the mean age in the subsample. Clearly, the women in this average and above-average group have experienced gains in life expectancy over the almost three decades covered by our data.

What has happened to the mortality differential between women in the low-family-earnings and high-family-earnings groups over that period? Figure 14 shows the mortality rate ratio of the low- and the high-family-earnings groups. Each highlighted dot along the lines shows the central point estimate of this ratio in a birth cohort and within an age group, with the mortality rate ratio calculated at the mean age in the subsample. Because the ratios of the younger birth cohorts are above those of the older cohorts, we can infer that mortality rate differences have grown over time. The parameter that captures the growing mortality difference is the coefficient β_4, and this coefficient is statistically significant for women in most age groups between ages 53-61 and 81-89 (the exceptions are ages 65-73 and 69-77). Thus, the results imply that mortality differentials have increased among women as they have among men, despite the fact that our indicator of married women's family earnings position is less precise than our measure of men's earnings position.

Status as indicated by respondents' relative educational attainment. The last results reported here show the impact of respondents' relative educational attainment on mortality differentials over time. Both men and women are classified by their educational attainment relative to that of other people who attained age 50 in the same year. Our classification distinguishes between men and women who have above-median and below-median educational attainment. When there are many ties at the 50th percentile of educational attainment, we randomly assign the people with tied rank to the higher or lower attainment group to attain the target percentage of people in the lower and higher education groups.[5] Even though this method of assigning respondents to schooling categories has shortcomings, the problems are less severe than those associated with classifying respondents from different generations to broad education categories, such as, "college graduate" or "less than a high school diploma."

Our statistical results suggest, not surprisingly, that respondents' relative educational rank has a significant impact on mortality. Both for men and women, respondents with below-median educational rank have higher mortality rates. Figure 15 shows the ratio of mortality in the low-

[5] An alternative procedure would have been to assign respondents an educational attainment score, which would have expressed respondents' attainment as a percent of the median school attainment of men or women in their birth cohort. Most schooling attainment levels can be expressed in terms of the number of years required to obtain them. One challenge to this method is that the Census Bureau changed its education codes in the early 1990s to emphasize degree attainment rather than years of schooling begun or completed. Some of the SIPP surveys use the old coding method; others use the new one.

schooling group relative to the high-schooling group for men in four birth cohorts. The results were obtained by dividing respondents' observations into overlapping 7-year age groups and then estimating equation 1 using the person-year observations that fall within the 7-year age spans (49-55, 52-58, 55-61, 58-64, 61-67, etc.). Each highlighted dot along the lines shows the central point estimate of the ratio of low-education mortality to high-education mortality in a birth cohort and within an age group, with the mortality rate ratio calculated at the mean age in the subsample. The mortality rate ratio has clearly been increasing at all ages. For each birth cohort, the age-specific death rate is higher for low education respondents, and the difference in mortality between low- and high-education respondents has widened in successive birth cohorts. The widening in mortality differences is highly statistically significant for all the male age groups between ages 58-64 and 73-79.

Figure 16 shows similar estimates for women. Again, results are obtained by dividing respondents' observations into overlapping 7-year age groups and then estimating equation 1 using the person-year observations falling in the 7-year age spans. As was the case for men, women with below-median schooling have higher age-specific mortality rates than women with higher levels of school attainment. It is less clear for women that the mortality differential has increased at every age. The estimated increase in the mortality differential is statistically significant for women in all but one of the age groups between 64-70 and 79-85. (The sole exception is the age group between 70 and 76.) For younger and older age groups the mortality differential is significant, but the increase in the differential is not.

Future research

This paper presents initial findings of a project on disparities in well-being among aged families. It examines the growth in income inequality, by age, in the period since 1979, and it increases our understanding of the growth in mortality differences between people with low and high career incomes. The ultimate goals of the project are, first, to determine the relationship, if any, between the delay in the average retirement age and the increase in inequality among people who are near retirement age; and, second, to estimate the change in the Social Security benefit formula that may be needed to compensate low-lifetime-earnings workers for the increases in mortality differences that have been observed in Social Security earnings and mortality records.

Refernences

Bosworth, Barry P., and Kathleen Burke. 2012. "Changing Sources of Income among the Aged Population." CRR WP 2012-27. (Chestnut Hill, MA: Center for Retirement Research at Boston College).

Bosworth, Barry P., and Gary Burtless. 2010. "Recessions, Wealth Destruction, and the Timing of Retirement." CRR WP 2010-22. (Chestnut Hill, MA: Retirement Research Center at Boston College).

Bosworth, Barry P., Gary Burtless, and Kan Zhang. 2014. "Sources of Increasing Differential Mortality among the Aged by Socioeconomic Status." Paper presented at the 16th annual joint meeting of the Retirement Research Consortium, August 7-8, 2014, Washington, D.C.

Burtless, Gary. 2013. "Who Is Delaying Retirement? Analyzing the Increase in Employment at Older Ages," in Henry Aaron and Gary Burtless (eds). *Closing the Deficit: How Much Can Later Retirement Help?* (Washington, DC: Brookings).

Compson, Michael. 2011. "The 2006 Earnings Public-Use Microdata File: An Introduction." *Social Security Bulletin* 71, no. 4: 33-59.

DeNavas-Walt, Carmen, and Bernadette D. Proctor. 2014. *Income and Poverty in the United States: 2013*. U.S. Census Bureau, Current Population Reports, P60-249. (Washington, DC: U.S. Government Printing Office).

Larrimore, Jeff, Richard V. Burkhauser, Shuaizhang Feng, and Laura Zayatz. 2008. "Consistent Cell Means for Topcoded Incomes in the Public Use March CPS (1976–2007)," *Journal of Economic and Social Measurement 33*(2): 89-128.

Meara, Ellen, Seth Richards, and David M. Cutler. 2008. "The Gap Gets Bigger: Changes In Mortality and Life Expectancy, By Education, 1981-2000." *Health Affairs* 27: 350-360.

Olshansky, S. Jay, and others. 2012. "Differences In Life Expectancy Due To Race And Educational Differences Are Widening, And Many May Not Catch Up." *Health Affairs*, 31, no. 9:1803-1818.

Singh, Gopal K., and Mohammad Siahpush. 2006. "Widening Socioeconomic Inequalities in US Life Expectancy, 1980–2000," *International Journal of Epidemiology* 35:969–979.

U.S. Congressional Budget Office. 2013. *The Distribution of Household Income and Federal Taxes, 2010.* (Washington, DC: U.S. Congressional Budget Office).

Waldron, Hilary. 2007. "Trends in Mortality Differentials and Life Expectancy for Male Social Security-Covered Workers, by Socioeconomic Status." *Social Security Bulletin* 67, no. 3:1-28.

Table 1. The Matched SIPP-Social Security Sample

	Male	Female	Total
Respondents with matched earnings and death records	**70,358**	**81,790**	**152,148**
Respondents with nonzero midcareer earnings	**54,558**	**56,184**	**110,742**
Respondents who are "married, spouse present"			**100,974**
Successfully matched to spouse in household			98,674
Matched households where both spouses have SS earnings files			60,418
Unique respondents used in mortality regressions based on earnings	**54,558**	**37,121**	
Person-year observations in mortality regressions based on earnings	682,650	507,108	
Total Deaths (up to 2012, between ages 49-91)	15,103	6,600	
Unique respondents used in mortality regressions based on education	**70,358**	**81,790**	
Person-year observations in mortality regressions based on earnings	710,001	856,501	
Total Deaths (up to 2012, between ages 49-91)	16,113	15,353	

Figure 1.Trends in the Gini Coefficient of Money Income Inequality, All Persons Regardless of Age of Family Head, 1979-2012

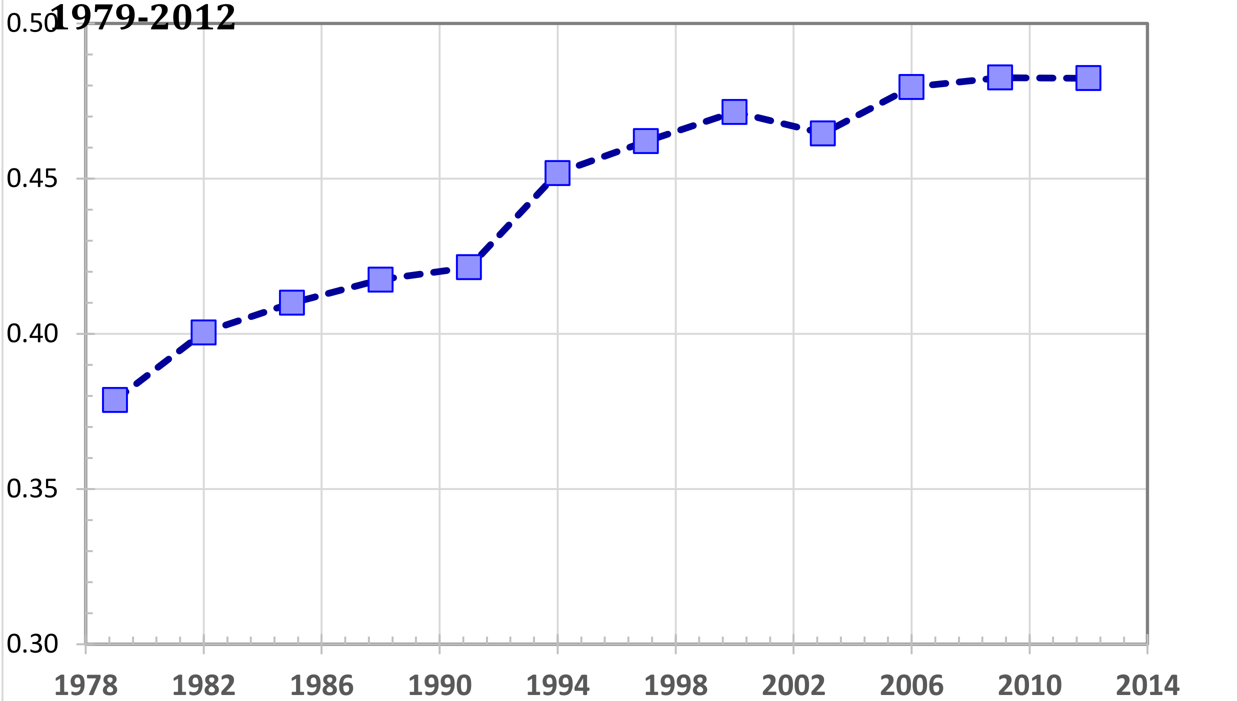
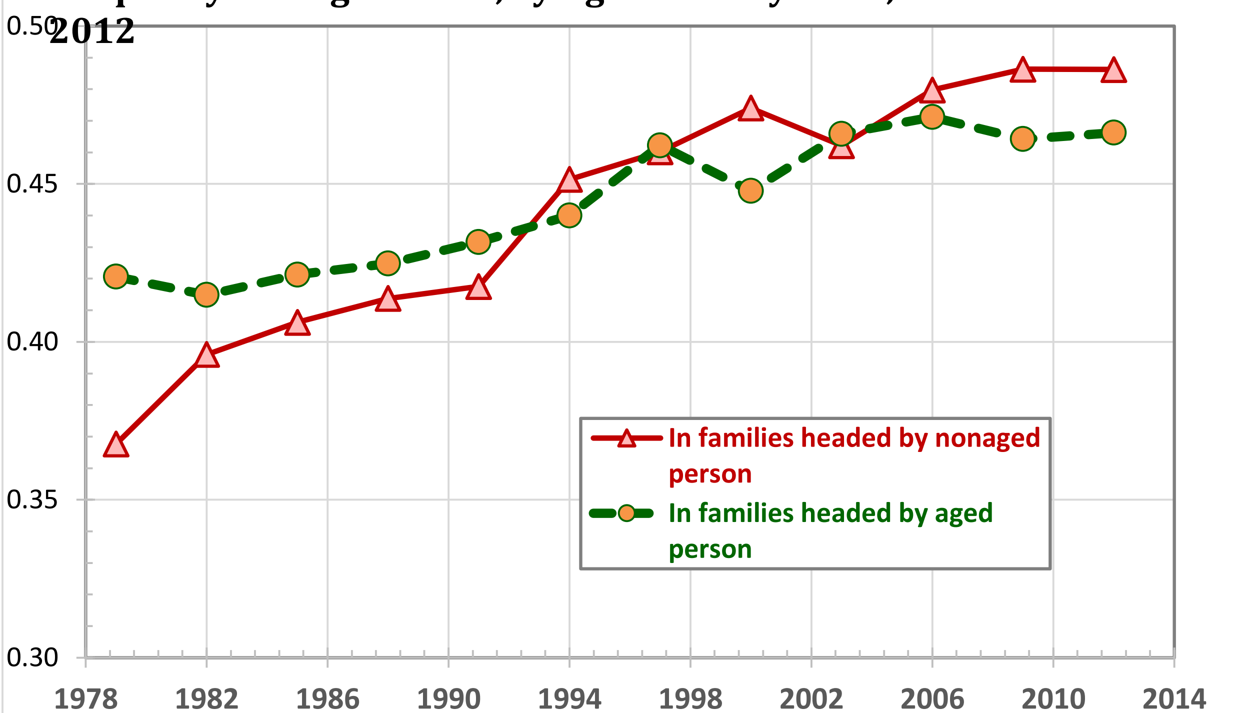


Figure 2. Trends in the Gini Coefficient of Money Income Inequality among Persons, by Age of Family Head, 1979-2012



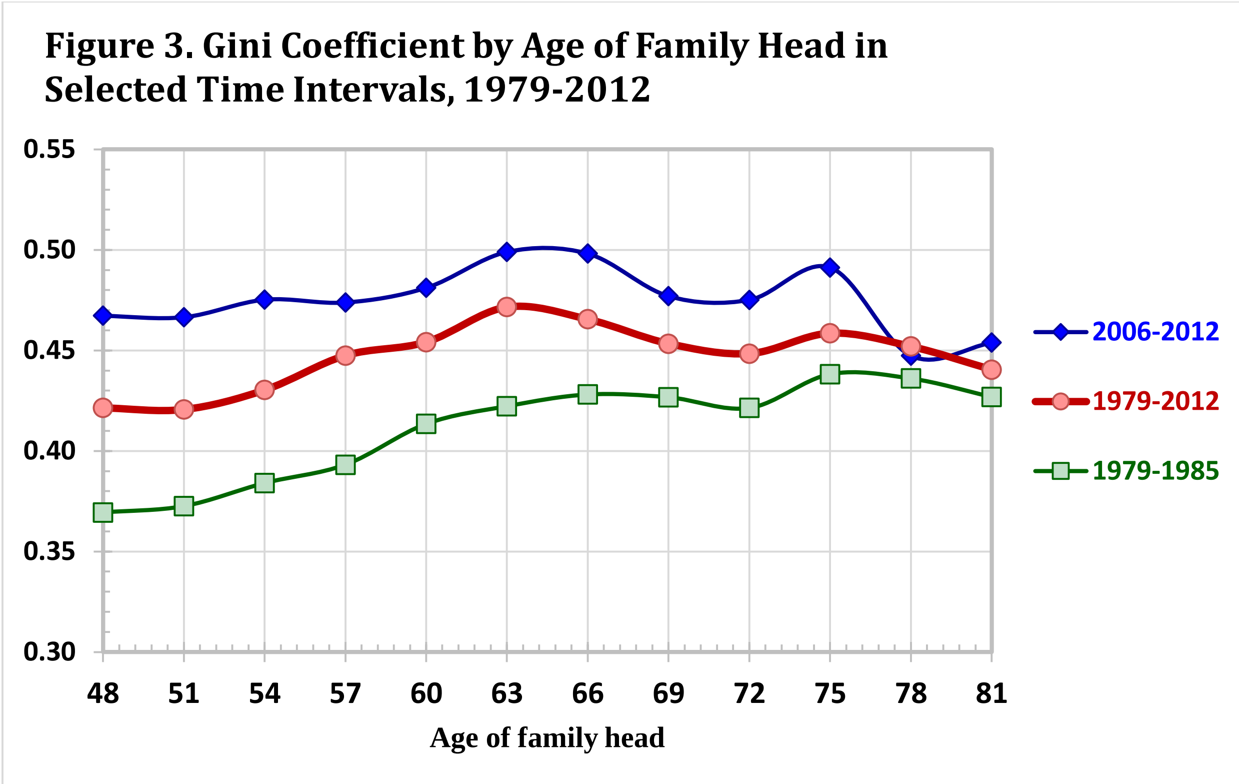


Figure 3. Gini Coefficient by Age of Family Head in Selected Time Intervals, 1979-2012
0.55
0.50
0.45
0.40
0.35
0.30
48
51
54
57
60
63
66
69
72
75
78
81
Age of family head
2006-2012
1979-2012
1979-1985

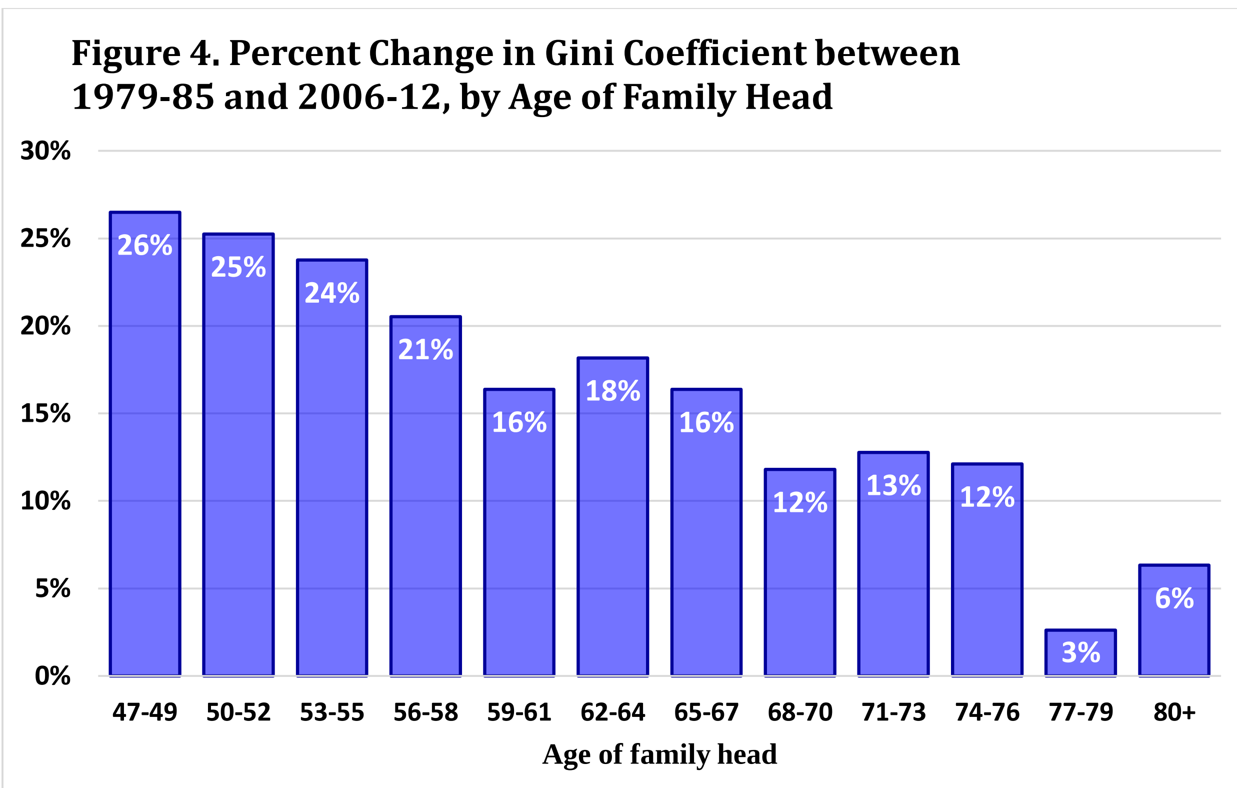


Figure 4. Percent Change in Gini Coefficient between 1979-85 and 2006-12, by Age of Family Head
30%
25%
20%
15%
10%
5%
0%
26%
25%
24%
21%
16%
18%
16%
12%
13%
12%
3%
6%
47-49
50-52
53-55
56-58
59-61
62-64
65-67
68-70
71-73
74-76
77-79
80+
Age of family head

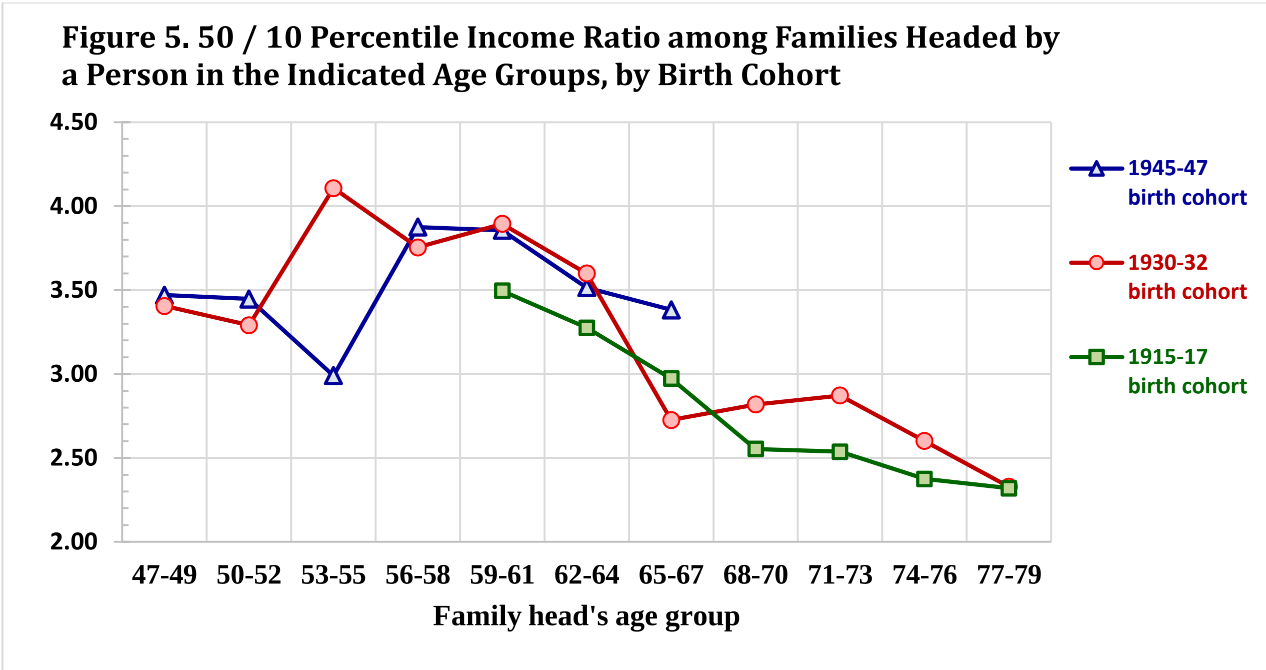


Figure 5. 50 / 10 Percentile Income Ratio among Families Headed by a Person in the Indicated Age Groups, by Birth Cohort
4.50
4.00
3.50
3.00
2.50
2.00
47-49
50-52
53-55
56-58
59-61
62-64
65-67
68-70
71-73
74-76
77-79
Family head's age group
1945-47 birth cohort
1930-32 birth cohort
1915-17 birth cohort

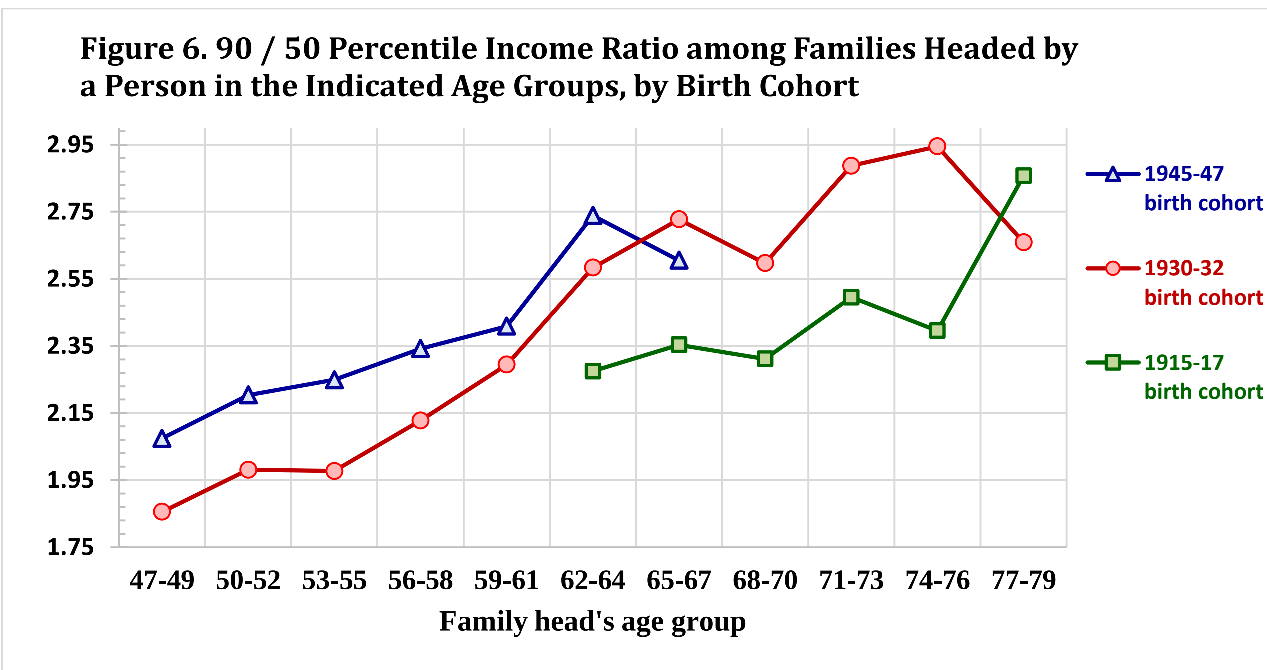


Figure 6. 90 / 50 Percentile Income Ratio among Families Headed by a Person in the Indicated Age Groups, by Birth Cohort
2.95
2.75
2.55
2.35
2.15
1.95
1.75
47-49
50-52
53-55
56-58
59-61
62-64
65-67
68-70
71-73
74-76
77-79
Family head's age group
1945-47 birth cohort
1930-32 birth cohort
1915-17 birth cohort

Figure 7a. Probability of Death among Males Born in 1925, by Age and Average Earnings between Ages 41 and 50 (Low earners below 31st percentile earnings in at least half of ages 41-50)

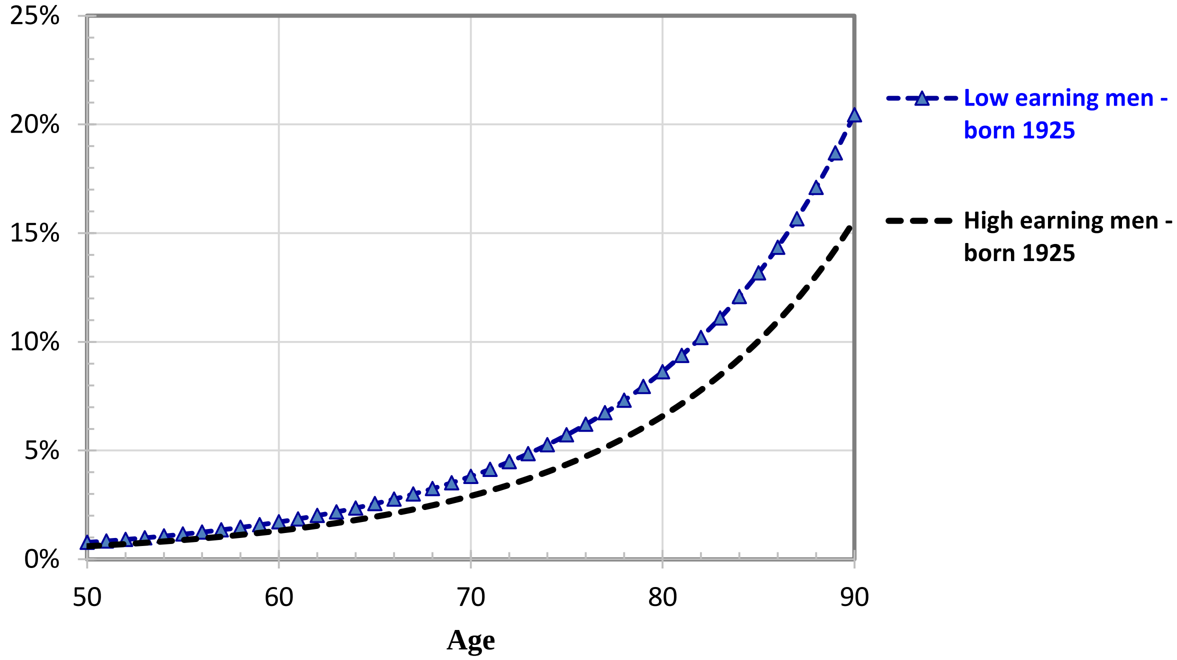


Figure 7b. Probability of Death among Males Born in 1945, by Age and Average Earnings between Ages 41 and 50 (Low earners below 31st percentile earnings in at least half of ages 41-50)

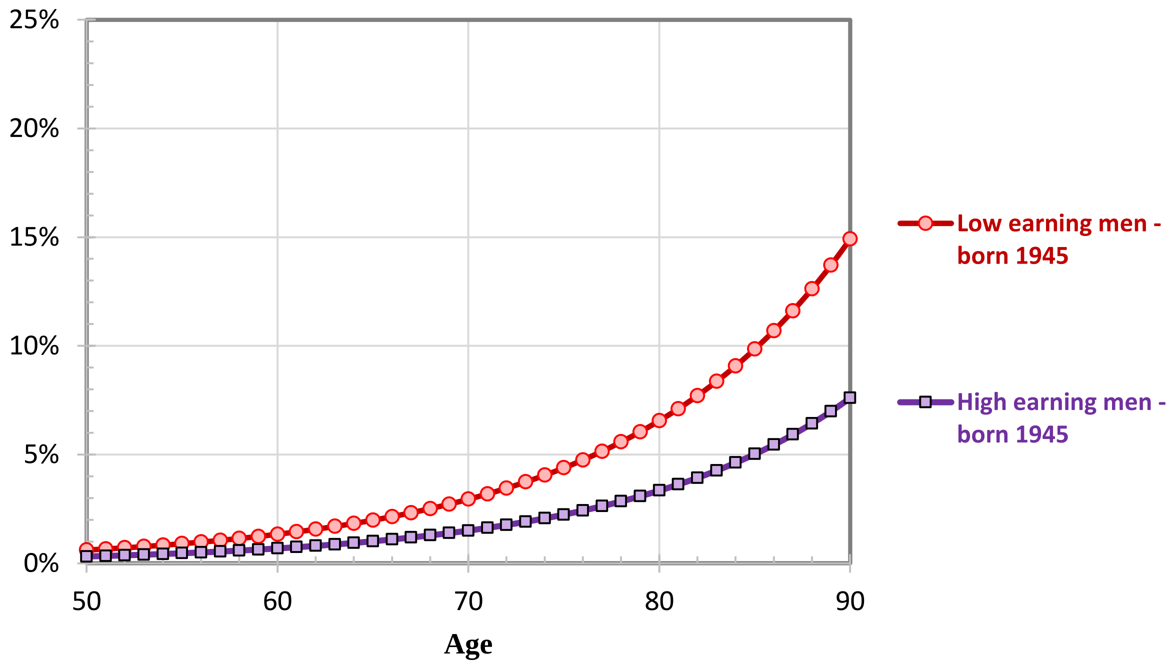


Figure 8. Mortality Rate of Average and High Earning Men Born in Selected Years, by Age (Low earners below 31st percentile earnings in at least half of ages 41-50)

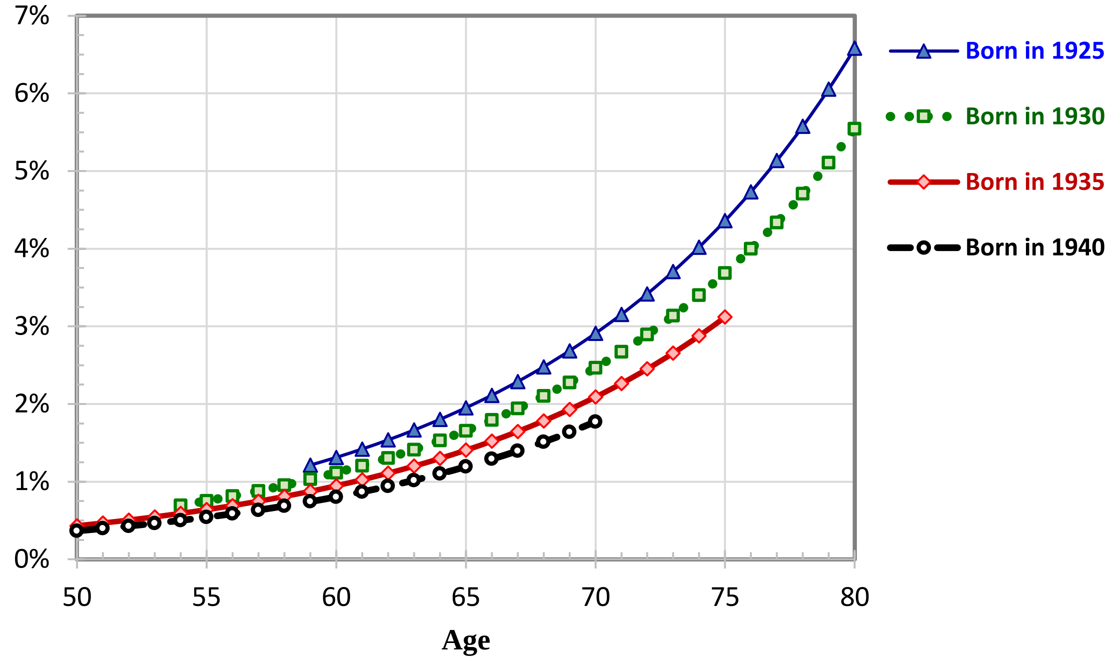


Figure 9. Ratio of Mortality among Low Earning Males Compared with Average and High Earning Males (Low earners below 31st percentile earnings in at least half of ages 41-50)

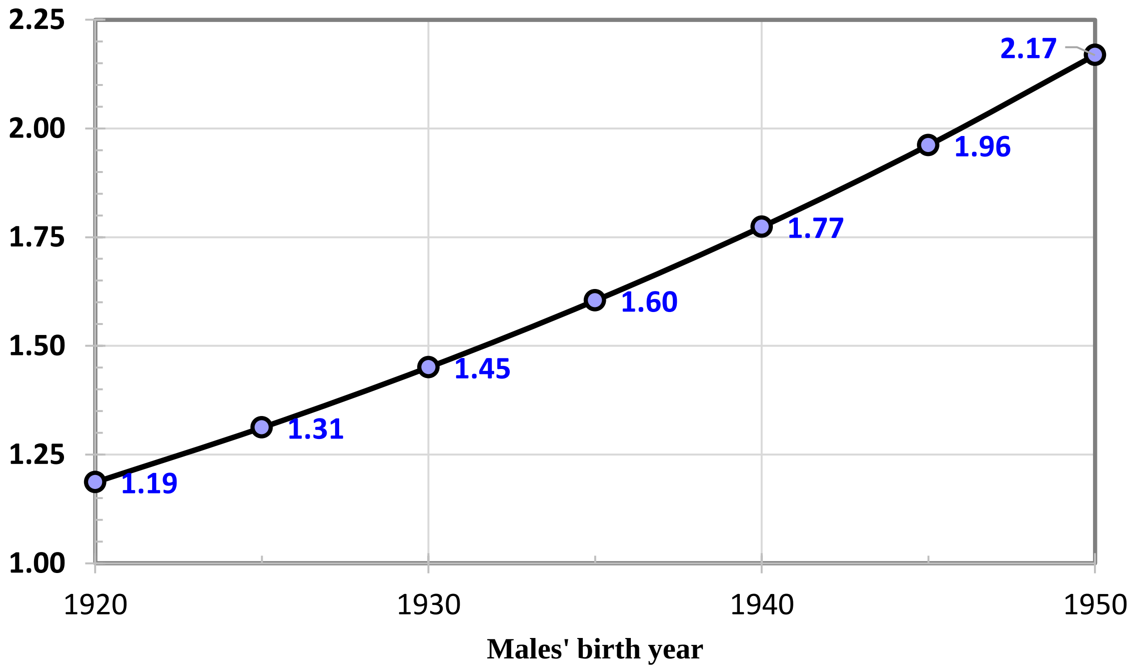
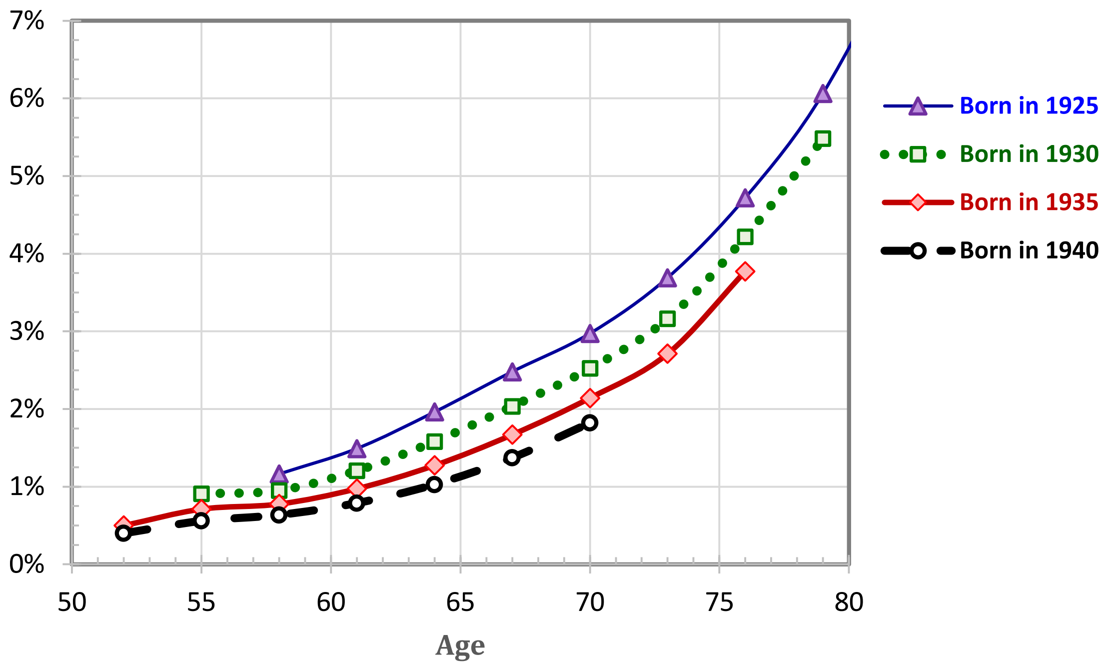


Figure 10. Mortality of Average and High Earning Men Born in Selected Years, by Age (Low earners below 31st percentile earnings in at least half of ages 41-50)



Fiugre 11. Mortality Rate Ratios of Low Earning Men Compared with High Earning Men Born in Selected Years, by Age (Low earners below 31st percentile earnings in at least half of ages 41-50)

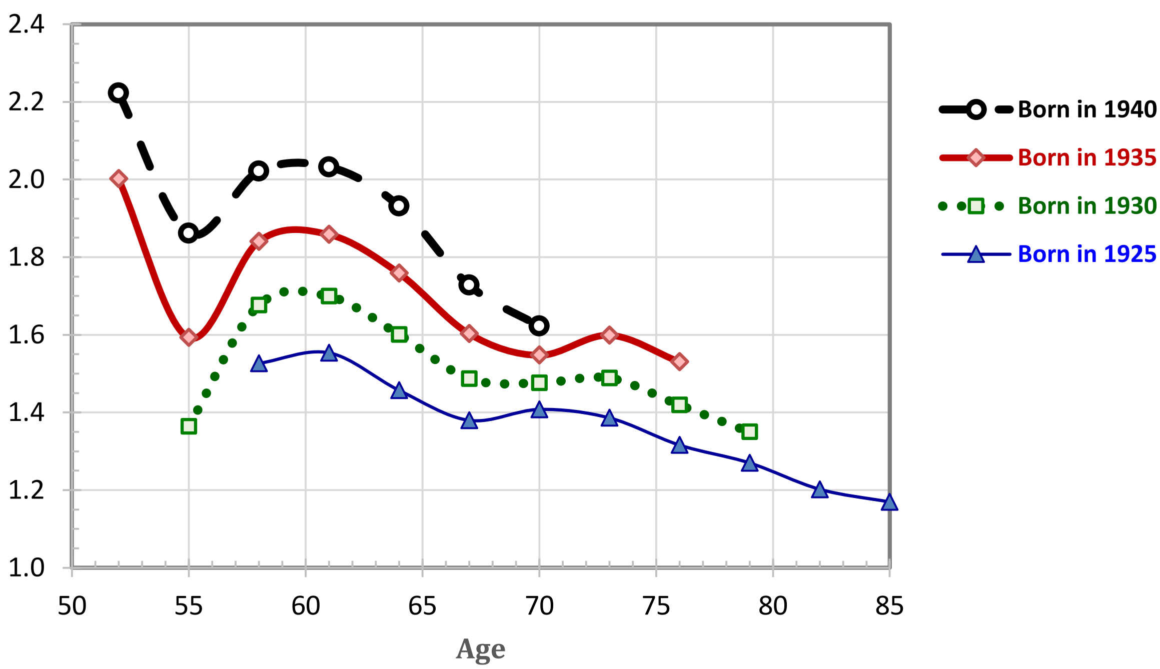


Figure 12. Excess Mortality of Low Income Women Born in Selected Years, by Age (Own earnings below 67th percentile of female earnings or spouse with low earnings)

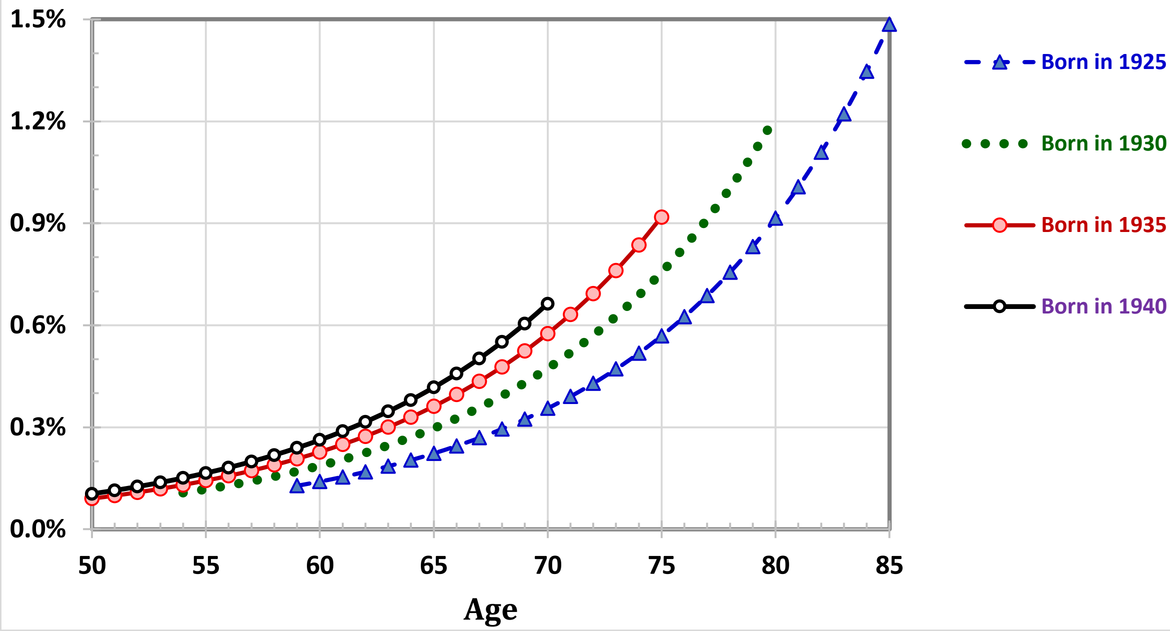


Figure 13. Mortality of Average and Above-Average Income Women Born in Selected Years, by Age (Wome have own earnings above 67th percentile or a average or high earning spouse)

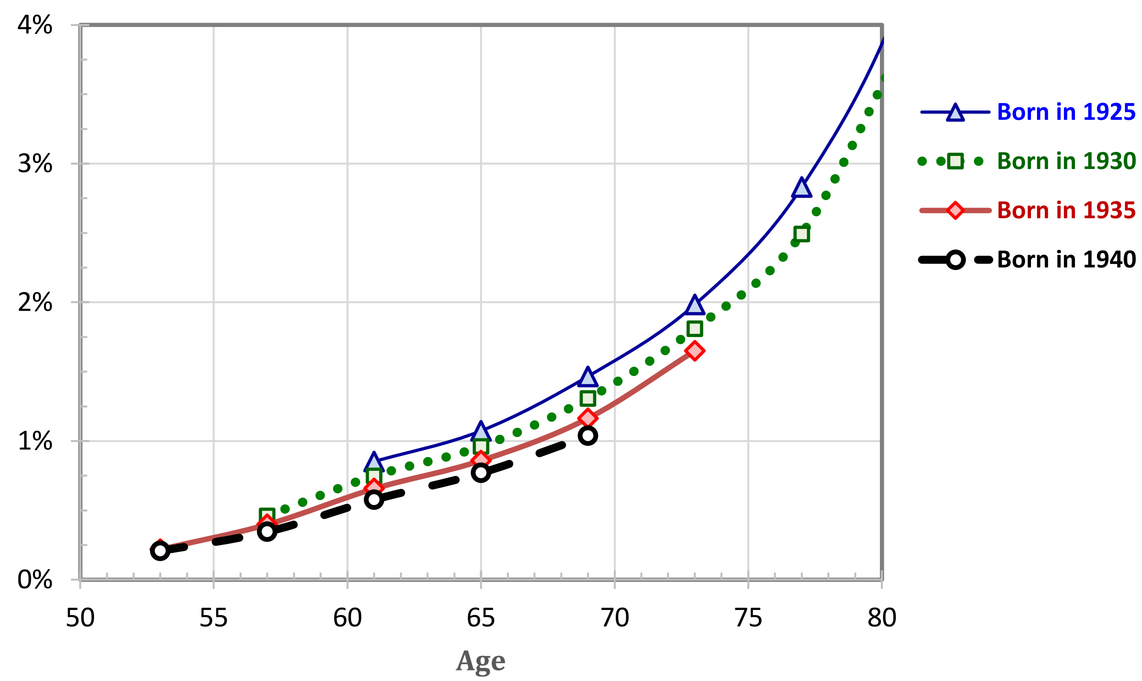


Figure 14. Mortality Rate Ratios of Low Income Women Compared with High Income Women Born in Selected Years, by Age (Low income have own earnings below 67th percentile or a low earning spouse)

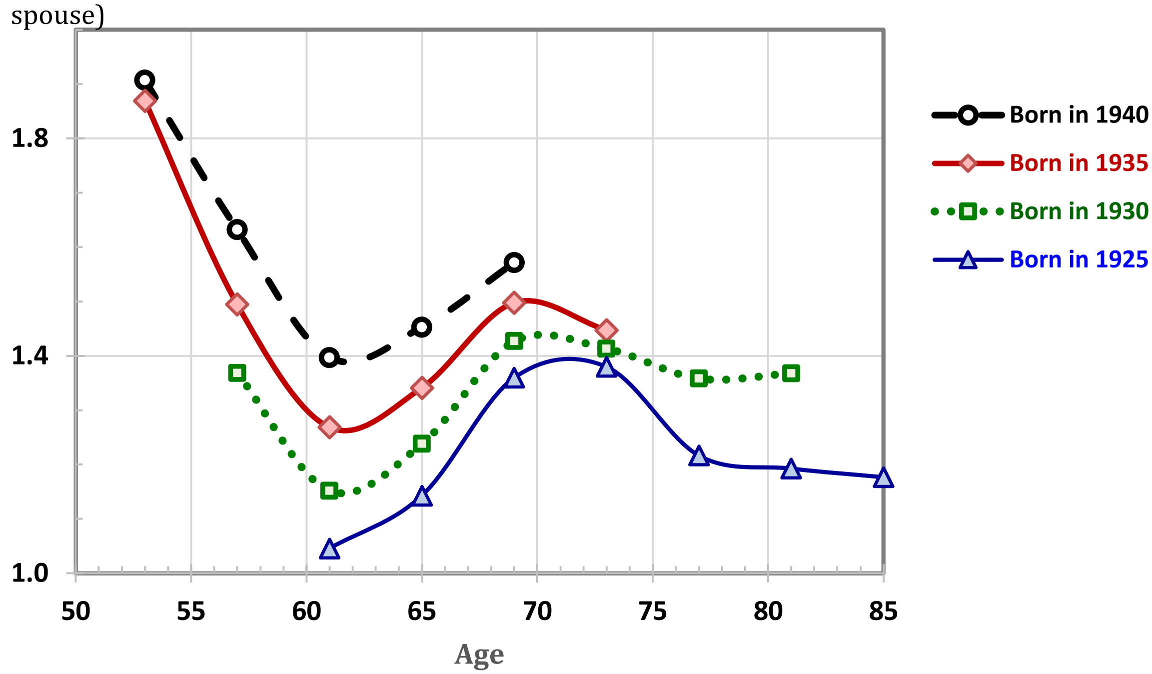


Figure 15. Mortality Rate Ratios of Low Education Men Compared with High Education Men Born in Selected Years, by Age (Low education men are in bottom half of cohort education distrib.)

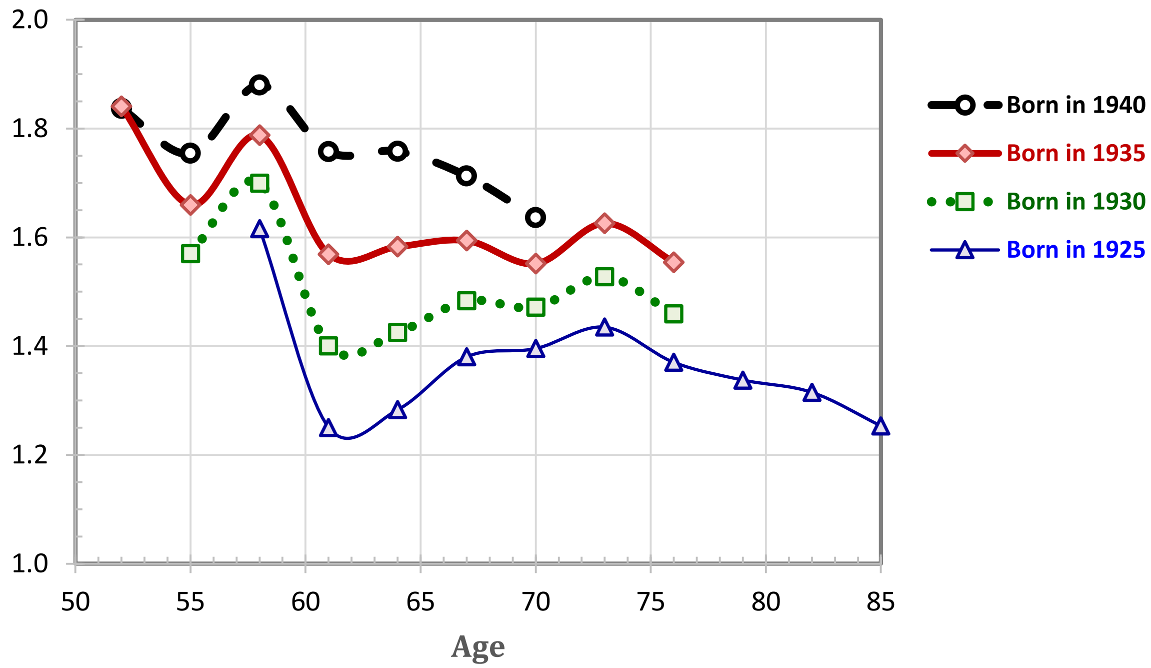


Figure 16. Mortality Rate Ratios of Low Education Women Compared with High Education Women Born in Selected Years, by Age (Low education women are in bottom half of cohort education distrib.)