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ABSTRACT

Social Security faces a major financing shortfall. One policy option for addressing this shortfall would be to raise the earliest age at which individuals can claim their retirement benefits. A welfare analysis of such a policy change depends critically on how it affects living standards. This paper estimates the impact of the Social Security early entitlement age on later-life elderly living standards by tracing birth cohorts of men who had access to different potential claiming ages. The focus is on the Social Security Amendments of 1961, which introduced age 62 as the early entitlement age (EEA) for retired-worker benefits for men. Based on data from the Social Security Administration and March 1968-2001 Current Population Surveys, reductions in the EEA in the long-run lowered the average claiming age by 1.4 years, which lowered Social Security income for male-headed families in retirement by 1.5% at the mean, 3% at the median, and 4% at the 25th percentile of the Social Security income distribution. The increase in early claiming was associated with a decrease in total income, but only at the bottom of the income distribution. There was a large associated rise in elderly poverty and income inequality; the introduction of early claiming raised the elderly poverty rate by about one percentage point. Finally, for the 1885-1916 cohorts, the implied elasticity of poverty with respect to Social Security income for male-headed families is 1.6 . Overall, we find that the introduction of early claiming was associated with a reduction in income and an increase in the poverty rate in old age for male-headed households.

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I. Introduction

The Social Security program is the largest single expenditure of the U.S. federal government, totaling $922 billion in 2016. It is also an important source of long-term fiscal imbalance. Due to the aging of the population, falling fertility, and a declining rate of productivity growth, Social Security currently faces a 75-year actuarial deficit equivalent to 2.83% of payroll (Board of Trustees, Federal Old-Age and Survivors Insurance and Federal Disability Insurance Trust Funds, 2017). For this reason, there are constant conversations among policy makers about reforms to the program. There is no shortage of ideas for making incremental changes, from raising the cap on payroll taxes, to investing some of the trust fund in stocks. More significant changes include raising the payroll tax rate or raising the Full Benefit Age (FBA) from its current level of 67. For each of these proposals, the welfare implications are fairly clear; the main debates are political.

A more radical change in the Social Security program would be to raise the Early Eligibility Age (EEA), currently 62. While age 67 is also known more colloquially as the "normal" retirement age in the U.S., there is nothing normal about it; by far the most common age for claiming Social Security benefits is 62, when retired-worker benefits first become available. So, an increase in the EEA would affect the timing and amount of benefits for a large number of older Americans. A number of OECD countries recently have increased the early entitlement age in their old-age pension systems, including Australia, Austria, Germany, and the United Kingdom, among others.

While we know that raising the EEA would reduce early retirement, we do not know the direction of the welfare effect, because the excess in claiming at age 62 in the U.S. is difficult to reconcile. As Diamond and Gruber (1999) emphasized, this spike cannot be explained easily by simple financial incentives, which are roughly actuarially fair for the population of individuals as a whole at that age. Rather, there are a number of competing explanations. The first is liquidity constraints: individuals would like to claim earlier than 62, but are unable to, because it is illegal to borrow against Social Security benefits. The second is that there is significant heterogeneity in mortality rates that make the financial incentives actuarially fair for the population as a whole, but actuarially unfair for some population sub-groups, who then rationally choose to claim early (Hurd et al., 2004). The third are behavioral failures: either due to misunderstanding (Brown et al., 2013), misoptimization (from, for example, cognitive limitations or financial illiteracy), or quasi-hyperbolic preferences (Diamond and Koszegi, 2003), individuals do not appreciate the fact that by taking benefits at age 62, they are lowering their monthly check until they die. The problem for welfare analysis is that these alternative explanations have directly opposite predictions for the welfare impacts of raising the EEA: the first suggests that it would lower welfare by tightening liquidity constraints; the second suggests that it could lower welfare for sub-groups of the population that had above-average mortality; while the third suggests that it would raise welfare by decreasing “mistakes.”

There is no study that comprehensively evaluates these tradeoffs, and we will not either. But what we offer in this paper is a piece of the puzzle: a demonstration of the impact on long-run living standards of early benefit claiming. If individuals who claimed benefits early did not see a diminishment of their later-life living standards (e.g. through other sources of income or smoothing through savings (Milligan, 2014; Bronshtein et al., 2018)), then it would suggest little welfare cost to allowing them to claim benefits earlier. Conversely, if early claiming is associated with reduced living standards, it raises a potential tradeoff from early claiming. This is especially true if the reduced living standards are among those with the lowest incomes.

This problem cannot be addressed by simply comparing those who do and do not claim Social Security benefits early, as these groups may differ in many ways that impact their later living circumstances. What is needed is an exogenous shift in the availability of early retirement options.

The introduction of the EEA into the Social Security program in 1961 provides that exogenous variation. Men born in 1896 and earlier moved through their early 60s prior to the 1961 Amendments to the Social Security Act. They could first claim benefits at age 65, the full benefits age (FBA). Men born in 1897 and later were affected by the law change. Those born in 1897 could claim as early as age 64; those born in 1898 could claim as early as 63; those born in 1899 and later could claim as early as age 62. With early claiming, age-65 benefits were actuarially reduced by 5/9ths of a percent for each month of payable benefits before attainment of age 65. For a man claiming at age 62, this represented a 20% reduction in benefits, compared to claiming at age 65. The cohort variation in the availability of early eligibility allows us to analyze its impact on financial well-being while controlling for both age and time trends.

The paper begins with a description of the 1961 Amendments and other relevant changes to the Social Security system around that time, then moves on to the empirical analysis, which focuses on men born in 1885 through 1916, a roughly fifteen year-of-birth window around the pivotal birth cohorts affected by the 1961 Amendments. These men attained their FBA and made their retirement decisions in the 1950s, 1960s, and 1970s. We then use data from the 1968-2001 March Current Population Surveys (CPS) to measure each cohort's income and poverty trajectories into old age. We use a regression framework to estimate the impact of the reduction in the early entitlement age on income and poverty, controlling for both age and calendar-year effects, as well as a broad set of potentially confounding factors.

There are three primary findings. First, reductions in the early entitlement age in the long-run lowered the average claiming age by 1.4 years, which lowered Social Security income for male-headed families in retirement by 1.5% at the mean, 3% at the median, and 4% at the 25th percentile of the Social Security income distribution. Second, the increase in early claiming was associated with a decrease in total income – but only for the lower half of the income distribution. As a result, there was a sizeable increase in elderly poverty and income inequality. In particular, the introduction of early claiming raised the elderly poverty rate by about one percentage point. Finally, for the 1885-1916 cohorts, the implied elasticity of poverty with respect to Social Security income for male-headed families is -1.6. Overall, we find that the introduction of early claiming was associated with a reduction in income and an increase in the poverty rate in old age for male-headed households.

These findings do not prove that there were negative welfare consequences from introducing the EEA. In particular, individuals may have been making a rational decision to trade off leisure for income. But they do rule out the notion early claiming was a clear welfare improvement and motivate further analysis of the welfare consequences of this key policy parameter.

The paper is organized as follows. The next section briefly relates our analysis to the previous literature. Then section III gives background on the 1961 Amendments. Section IV describes the CPS data and discusses cohort trends in income and poverty, whereas Section V lays out the regression framework. Both of these sections draw on the organization, exposition, and methodology developed in a companion set of papers, Engelhardt and Gruber (2006), Engelhardt, Gruber, and Perry (2005), and Engelhardt (2008). Section VI presents the empirical results. There is a brief conclusion.

II. Previous Literature

Our analysis is most closely related to five strands of the empirical literature on Social Security. The first is a large, longstanding literature that has examined the role of Social Security in retirement in the 1960s and 1970s. Reviewed in detail in Feldstein and Liebman (2001), it includes notable contributions by Crawford and Lilien (1981), Moffitt (1987), Rust and Phelan (1997), among many others. This body of work generally concluded that decreasing the age of eligibility to 62 for men not only increased Social Security claims, but also led to earlier departures from the labor force, with a spike emerging in the retirement hazard at age 62. Related work in other contexts includes the option-value approach of Stock and Wise (1990) to analyze the timing of claims to private pensions, and Gruber and Wise (1999), who documented similar evidence on the impact of early claiming in Social Security programs in a wide variety of OECD countries.

The second strand is comparatively more recent and has used both structural models and reduced-form approaches to examine the impact of early and delayed claiming on labor supply, saving, and retirement income adequacy. This includes work by Blau (2008), Gustman and Steinmeier (1986, 2008, 2015), Coile et al. (2002), Sass et al. (2014), Shoven and Slavov (2014), and Shoven et al. (2017, 2018), among many others.

The third strand has examined the short- and long-run impact of benefit increases and cuts, respectively, from the 1971 and 1977 Social Security Amendments that generated the so-called Social Security "notch." These amendments changed real benefits, while leaving the FBA and EEA unchanged at (then) 65 and 62, respectively. This includes work by Krueger and Pischke (1992), who found little evidence of an impact on labor supply, and Engelhardt and Gruber (2006), Engelhardt, Gruber, and Perry (2005), and Engelhardt (2008), who found substantive long-run impacts on old-age poverty, shared living arrangements, and home ownership.

A fourth, more recent strand has estimated the short-run impact of the 1983 Social Security Amendments that increased the FBA from 65 to 67, while leaving the EEA intact at 62, which effectively generated benefit cuts. This includes work by Mastrobuoni (2009) and Behagel and Blau (2012), who examined the impact on labor supply and claiming behavior, and Duggan et al. (2009), who examined the spillover effects into the federal disability insurance (DI) program.

The final strand consists of a set of a current empirical studies that examine the short-run impacts of recent increases in the EEA that varied by birth cohort on labor force participation, poverty, and program participation of older individuals in other countries. This includes Atalay and Barrett (2015), who studied women in Australia, Staubli and Zweimüller (2013) and Manoli and Weber (2016), who studied men and women in Austria, and Cribb and Emmerson (2017), who studied women in the United Kingdom. Taken together, these studies show that for individuals close to the EEA (i.e., in their late 50s and early 60s) raising the EEA generates significant increases in participation in the labor force and other gateway social insurance programs (e.g., disability and unemployment insurance, with results differing somewhat across countries). Cribb and Emmerson (2017) showed that raising the EEA actually raised the poverty rate in the short-run, because the increases in earnings and other program income were not enough to offset the loss of old-age pension benefits.

Instead of focusing on the labor-market impact of early eligibility like these studies, our main contribution is to examine the long-term impact of early claiming on well-being, as measured by family income and poverty. We do so by examining the introduction of the EEA for men in the early 1960s, which differed by birth cohort, and then trace the trajectories of income and poverty as the affected cohorts progressed into old age.

III. Legislative History

The 1950s, 1960s, and 1970s were central to the formation of today's Social Security program.[1] This section describes the focal program changes in this period and illustrates their time-series impact on benefit awards in Figures 1-5, which are based on data drawn from various issues of the Social Security Administration's *Annual Statistical Supplement*.

In the early 1950s, age 65 was the earliest an individual could receive any type of benefit based off career earnings. At that point, someone could claim retired-worker benefits equal to 100% of the primary insurance amount (PIA) based on his/her covered earnings history. The 1956 Social Security Amendments granted women the opportunity to claim actuarially-reduced benefits as early as age 62 (Schottland, 1956). Early claiming lowered the age-65 benefit for women by 5/9ths of one percent for each month of payable benefits prior to age 65 and applied both to retired-worker and wife's benefits. For a woman claiming early on her 62nd birthday, this amounted to a

[1] Parsons (1991) and Ransom and Sutch (1986) have examined retirement behavior in the early years of the Social Security program. The legislative history for Social Security in this period is detailed in Cohen and Ball (1965, 1967), Cohen et al. (1966), Myers (1964), and Cohen and Mitchell (1961).

20% reduction in benefits. The 1956 Amendments took effect for calendar year 1957. The 1956 Amendments also introduced the disability insurance (DI) program, in which insured workers with a demonstrated need could draw benefits prior to age 65.

The Social Security Amendments of 1961 extended early claiming on the same basis to men (Cohen and Mitchell, 1961). Men born in 1897 and later were affected by the law change: those born in 1897 could claim as early as age 64; those born in 1898 could claim as early as 63; those born in 1899 and later could claim as early as age 62. Overall, the 1961 Amendments induced age-by-cohort variation in the eligibility for early claiming: men born in 1897-1899 became (partially or fully) eligible for early claiming in a manner that depended on age. Those born 1900 or later were eligible to claim early at all ages from 62-64. Those born prior to 1897 were not eligible for early claiming. This is the central variation that is used in the empirical analysis below.

Figure 1 illustrates the impact of these amendments on total retired-worker claims. The figure plots the aggregate number of claims for retired-worker benefits for both sexes pooled from 1955-1975[2] There are three spikes in the time series. The first occurred in 1957, when early claiming for women was introduced. The second occurred in 1961-1962, when early claiming for men was introduced and fully phased in. The third occurred in 1966, the year of the introduction of Medicare.[3]

[2] The number of insured workers grew smoothly in this period. Early issues of the *Annual Statistical Supplement* pooled these data for the sexes, which, unfortunately, means the series cannot be broken out for men only.

[3] The explanation for this third spike is somewhat complicated. Medicare was enacted in 1965, but was not implemented until July, 1966. The spike in 1966 occurred because many insured workers who had not yet claimed were required to claim OAS benefits in order to become eligible for Medicare benefits at that time (in the early years, Medicare was administered by SSA). This peculiar feature of the initial administration of the Medicare program is described in Ball (1966). Specifically, prior to 1965, there were a large number of individuals who were 65 and older, who were fully insured for OAS benefits, but did not claim, because of the retirement earnings test (RET). During this period, the earnings test applied to beneficiaries under age 72, and it was all-or-nothing: if a beneficiary earned

Figures 2 and 3 break this time series out by sex and age group. There are three notable features in Figure 2, which shows the number of new retired-worker awards by age group for women. First, there was an upward trend toward retirement at ages 65-66 in the early 1950s that reversed itself with the introduction of early eligibility in 1957. The same occurred for men in Figure 3: there was an upward trend toward retirement at ages 65-66 throughout all of the 1950s that reversed itself with the introduction of early eligibility in 1961. Together, the figures show a significant shift in the composition of claiming, away from claiming at the Full Benefit Age (FBA) of 65, toward early claiming at ages 62-64. Second, the increase in the number of claims at ages 62-64 (in 1957 for women and 1961 for men) exceeds the decline in claims at ages 65-66. Third, there is a secular decline in claiming at ages 67 and older after 1957 for both sexes, and, although interrupted in 1966 when Medicare came online, this decline continued thereafter. After the introduction of Medicare, claiming retired-worker benefits at 67 or older was relatively uncommon.

The dashed line in Figure 4 shows the time series of the percent of new awards for men that were due to early claiming, as measured on the left-hand vertical axis. After the introduction of early claiming in 1961, this share rises across years. In 1975, about 55% of all new retired-worker claims were early claims. The solid line in the figure plots the time series of the real value of new awards for men, as measured on the right-hand vertical axis. While generally increasing

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$1 above the earnings test threshold, no benefits were paid. The benefit award was deemed "conditional and deferred," with the benefit reduction under the test returned through higher benefits in future years. However, as described in Ball (1966), this structure of the earnings test, in practice, had the effect of deterring claiming for individuals between the ages of 65 and 71. Specifically, among those fully insured for benefits, but who expected to earn above the earnings test threshold, many did not claim benefits, because those benefits would have been clawed back fully. In anticipation of this, many insured workers simply did not claim. Ball (1966) estimated that in 1965, over one million insured workers were not claiming because of the earnings test. When Medicare was enacted, these individuals claimed OAS benefits in order to get health insurance, generating the large spike in the figure.

over time, the trajectory flattens after the introduction of early claiming. Not surprisingly, the actuarial reduction in benefits generates a strong inverse relationship between the share of claims from early claiming and the amount of new awards.

Figure 5 illustrates the initial take-up of early claiming by plotting the change in the age distribution of claims for new retired-worker benefits for men across the focal cohorts affected by the 1961 Amendments. Men in the 1895-6 cohorts were not eligible for early claiming, and the majority claimed at the FBA of 65. The age distribution shifted systematically leftward toward early claiming with the 1897, 1898, and 1899 cohorts, eligible at ages 64, 63, and 62, respectively. For the 1900 cohort, about one-third of men claimed at the FBA of 65 and one-fifth at the EEA of 62.

IV. Data Construction and Year-of-Birth Trends

The analysis focuses on men born in 1885 through 1916, a roughly fifteen year-of-birth window around the pivotal cohorts affected by the 1961 Amendments. We do not include women in the analysis for three reasons. First, most married women from similar cohorts claimed benefits on their husband's earnings history. Second, for widows and divorcees, the CPS did not ask about the year of birth for ex- or deceased husbands, so we cannot assign these women to the correct EEA for their husband's birth cohort. Finally, for never-married women, who would have claimed on their own earnings histories, the CPS sample sizes were too small for reliable analysis.

One concern, which will become apparent in the description of the regression framework below, is that cohorts in the latter half of this sample window experienced rapidly rising generosity of benefits for workers claiming at all ages. In particular, prior to 1971, Congress adjusted benefits on an ad hoc basis to account for the impact of inflation. Because of persistent high inflation in the late 1960s, the 1971 Social Security Amendments sought to codify inflation adjustment into the calculation of benefits on an automatic basis. However, the 1971 law inadvertently introduced the double indexation of benefits by both allowing the bend points in the PIA formula to be indexed to price inflation and basing Average Monthly Earnings (AME), the lifetime earnings measure at that time, on nominal wages, which already had real wage and inflation components. With high-inflation in the 1970s, this led to a rapid increase in benefits for each subsequent retiring birth cohort (as can be seen in Figure 6). This pattern of rising benefits ended with the 1977 Amendments, which eliminated double indexation by changing the PIA formula and introducing Average Indexed Monthly Earnings (AIME). Since those born in 1916 already would have attained the early retirement age of 62 in 1978 when the law went into effect, the 1977 Amendments grandfathered all individuals born in 1916 and earlier under the old benefit structure. For those born in 1917-21, the so-called "notch" generation, a new, less generous benefit structure was phased in.

In order to hold the benefit structure constant, as much as possible, the Social Security "notch" cohorts that began with men born in 1917 are excluded from the sample. Then, a cohort-specific measure of general benefit generosity is used in the regression specification below to control directly for changes in benefits that are not due to the introduction of early claiming per se.

In addition to rising benefits prior to the notch, the 1966 Amendments to the Social Security Act narrowly targeted supplemented benefits for individuals 72 and older who were not fully insured for benefits. This provision was known as the special Age-72 Benefit (Schobel, 1983), and was paid to a comparatively small number of individuals in 1966, which rapidly declined in subsequent years. The number of quarters of coverage required for eligibility for this supplement differed by year of birth. In the empirical analysis below, we control for eligibility for this benefit.

Men born 1885-1916 attained their FBA and made their retirement decisions in the 1950s, 1960s, and 1970s. To follow these cohorts into old age, the empirical analysis focuses on data from the 1968-2001 March CPS on elderly male-headed families aggregated into age-by-calendar-year cells. The questions in the March CPS are about income earned in the previous calendar year, so that the income data from the 1968-2001 surveys can be used to measure income and poverty status in 1967-2000. By the end of the sample period in 2000, men born in 1916 would have been 84, and men from older cohorts would have been even older or have died completely.

The sample construction follows the same methodology as in Engelhardt, Gruber, and Perry (2005), Engelhardt and Gruber (2005, 2006), and Engelhardt (2008), and begins with the CPS microdata. There, an elderly "family" is defined as the male head age 66 or older, his spouse, and any of his children living with the family and under the age of 18. Over time, the CPS has provided more disaggregated questions on income sources, and, for some types of income, has changed the wording of questions. The most disaggregated income measures are used to construct the two key measures of income: family Social Security income and total income. The former is the sum of Social Security retirement income across all members in the family, then adjusted for family size by the OECD equivalence scale. The latter is the sum of income from all sources across all members in the family, then adjusted for family size by the OECD equivalence scale. This measure is used to determine whether the family is above or below the federal poverty threshold. Finally, since the central variation in Social Security benefits from the introduction of the EEA is by year of birth, then poverty and income measures are collapsed into age-by-calendar-year cells, which are the same as year-of-birth cells. Both income measures are then deflated into real 2001 dollars using the all-items Consumer Price Index (CPI). Table 1 shows sample means of the key outcome and explanatory variables used below.

The solid line in Figure 6 shows real family Social Security income by year of birth based on the CPS data. Social Security income rises rapidly for men born 1885-1896, then falls slightly for men 1897-1900, which coincides with the cohorts first eligible for early claiming under the 1961 Amendments. After 1900, Social Security income continues a rapid rise through 1916, the end of the sample cohorts. The dashed line in the figure shows family total income, which trends fairly smoothly upward across these cohorts. There is some suggestion of a slowdown in income growth for birth cohorts in the late 19th century, but the correlation appears fairly weak.

Figure 7 plots real family Social Security income and the percentage of male-headed elderly families with total income below the federal poverty threshold (for the appropriate family size), the standard measure of absolute poverty. The poverty rate declines sharply from 42 percent for men born in 1885 to 20 percent for men born in 1896. For the 1897-1900 cohorts, the poverty rate rises above that for the 1896 cohort; after 1900, the poverty rate falls gradually as Social Security income rises. In contrast to Figure 6, there is stronger visual evidence here for a relationship between the early claiming policy change and poverty.

Unfortunately, this measure of poverty has a number of well-known limitations. First, it holds constant the standard of living. Second, it adjusts for price inflation, but not for real wage growth. Finally, it does not measure the depth of absolute deprivation. Consequently, Figure 8 shows Social Security income and a relative measure of poverty: the percentage of male-headed elderly families with income less than 40% of the median income of non-elderly families. Both elderly and non-elderly income are adjusted by the OECD equivalence scale. Non-elderly families are defined as those headed by someone 25-54 years old. The measure is constructed for elderly families by single year of age in each calendar year of the CPS, then collapsed to get a year-of-birth average, which is plotted in the figure. As Figure 8 shows, the results for the relative poverty rate are very similar to those for the absolute poverty rate.

Figure 9 shows Social Security income and a measure of income inequality for male-headed elderly families by year of birth. This measure is the 90-10 coefficient of variation. It is calculated by single year of age in each calendar year of CPS data as the difference in the 90th and 10th elderly OECD-equivalent income percentiles normalized by mean elderly income, and then collapsed to get a year-of-birth average, which is plotted in the figure. Elderly income inequality varies across years of birth and appears more weakly correlated with Social Security income. The sample correlation coefficient between the series is -0.33.

V. Regression Framework

Overall, these year-of-birth figures suggest that cohorts of men that were first eligible for early claiming had lower actual Social Security income and higher poverty rates after retirement and into old age. A fundamental challenge in interpreting these patterns as causal is that most of the variation in Social Security benefits that identifies differences in elderly family Social Security income across years of birth is time-series in nature. Omitted variables that are correlated with changes in poverty rates and Social Security, and trending over time, might explain these patterns equally well, leading to a fundamental identification problem. For example, lifetime earnings are affected by aggregate productivity and human capital accumulation that have changed across time and cohorts. As a key determinant of Social Security benefits, these changes would find their way into observed Social Security income. At the same time, federal poverty thresholds are inflation-adjusted, but not average-earnings adjusted. So, poverty rates would have been expected to have declined for successive birth cohorts as productivity, human capital accumulation, and real lifetime earnings rose. Hence, gains in productivity and human capital could simultaneously account for a rise in Social Security income and a decline in poverty across years of birth.

To attempt to circumvent this and identify causal impacts on poverty, the analysis moves to the following regression framework. Let a index age of the male head and t index the calendar year, then the econometric specification can be written as

$$S_{at} = \alpha + \beta EEA_{at} + \boldsymbol{\xi}_a + \boldsymbol{\psi}_t + \boldsymbol{\delta}'\mathbf{X}_{at} + f(Z_{at}^{\overline{65}}) + u_{at} . \tag{1}$$

The dependent variable S is the age-by-calendar-year mean family Social Security income; u is an error term. The focal explanatory variable is EEA, the early entitlement age for men in that cohort: for men born in 1896 or earlier, the EEA was 65; for men born in 1897, the EEA was 64; for men born in 1898, the EEA was 63; for men born in 1899 and later, the EEA was 62. $\boldsymbol{\xi}$ is a vector of dummy variables for single year of age; $\boldsymbol{\psi}$ is a vector of dummy variables for calendar year. The age dummies control for differences across age groups in the outcome measure; the year dummies control for any general time trends in the outcome measure. The central objective is to get a consistent estimate of β, the impact of a one-year increase in the EEA on Social Security income. Controlling for age and calendar year, the estimate of β is identified by cross-cohort variation in eligibility for early claiming induced by the 1961 Amendments.

The vector $\mathbf{X}$ includes controls for cell means of educational attainment of the head (high school diploma, some college, and college or advanced degree), marital status (married, widowed, and divorced), white, and veteran status. These account for any other trends in cohort characteristics that might be correlated with both the legislative changes in benefits determination and actual income. In addition, there are controls for important changes that varied by cohort: a dummy if men in that year of birth were eligible for Medicare at age 65; a dummy if men in that year of birth were eligible for DI prior to age 65; and the quarters of coverage required for that cohort to qualify for the special Age-72 supplemental Social Security benefit.

As Figure 6 indicated, Social Security income was trending upward across the sample cohorts. So as to prevent these more general benefit increases from confounding the estimates of β, (1) includes a polynomial function (f) of $Z^{\overline{65}}$, a measure of the simulated real family Social Security income a man would have gotten had he claimed benefits at age 65. For a synthetic unmarried male beneficiary from a given cohort, $Z^{\overline{65}}$ is constructed as follows. Let $B_c(y_c,k)$ be the primary insurance amount (PIA), which varies according to the benefit structure B_c applied to year of birth c, potential claiming age k, and earnings history y. First, an earnings history y was constructed for each cohort c. A baseline earnings history for the 1916 birth cohort—the last year of birth in the sample—was constructed from various issues of the Social Security Administration's *Annual Statistical Supplement*. In particular, SSA published median male earnings by five-year age groups on an annual basis. These earnings were used to assign median earnings at age 22 (from the median earnings for ages 20-24 in 1938), age 27 (from median earnings for ages 25-29 in 1943), etc., in five-year intervals. Then a linear trend in earnings was assumed within these five-year intervals to get earnings by single year of age for the 1916 cohort.

This method was used through age 60, and earnings were assumed to grow with inflation for ages 60-69.[4] Importantly, the earnings history constructed for the 1916 cohort, denoted as $\bar{y}_{1916}$, was then used for *all* birth cohorts, and the CPI was simply used to adjust the earnings profile for inflation for earlier and later cohorts. Therefore, all cohorts were assigned the same real earnings trajectory by construction.

Second, the constructed earnings histories for each cohort were input into the SSA's ANYPIA benefit calculator, which calculates the PIA at retirement, given a date of birth, date of retirement, and earnings history. In this case, ANYPIA calculates $B_c(\bar{y}_{1916}, 65)$. Unfortunately, the actual day and month of birthdates are not observed in the CPS. Because the timing of cost-of-living adjustments has varied across calendar years, for the purposes of calculating simulated benefits, birthdays of June 2 were assigned in the particular year of birth, and it is assumed that men retired and claimed benefits in June of the year in which age 65 is attained. Under this assumption about birthdays, men born in 1897 could have claimed at age 64; men born in 1898 could have claimed at age 63; men born in 1899 and later could have claimed at age 62. However, men born 1896 or earlier could have claimed only as early as age 65.

These PIAs, expressed in real 2001 dollars, are shown as the solid line in Figure 10. The variation in PIA at age 65, even conditional on constant earnings histories, is readily apparent in the figure. The PIA rises for early cohorts, then is roughly constant in real terms, followed by a rapid rise for those born after 1903 due to ad hoc benefit adjustments, then subsequently ramps up quickly through 1916 due to double indexation.

[4] Median earnings for workers over 60 are not used, because, by that age, many of the workers in the cohort have entered "bridge" jobs, and the median worker's earnings at these ages may not be representative of workers who have remained in their lifetime jobs through age 65.

Finally, the Social Security Administration periodically increased nominal benefits to adjust for inflation. To obtain a value for the PIA, which is measured at the time of retirement, in a future calendar year in which a cohort is observed in the CPS sample, all "cost of living adjustments" (COLAs) to which a beneficiary was entitled from the time of retirement until the date of interview, $\prod$, were incorporated to produce a real, expected Social Security *monthly* benefit. This was multiplied by 12 to convert it to an annual income amount and then adjusted by the OECD equivalence scale θ to yield

$$Z_{at}^{\overline{65},Unmarried} = 12 \cdot B_c(\overline{y}_{1916}, 65) \cdot \prod\nolimits_{act} \cdot \theta \ . \tag{2}$$

A synthetic married man from a given cohort was assigned 150 percent of this measure,

$$Z_{at}^{\overline{65},Married} = 1.5 \cdot [12 \cdot B_c(\overline{y}_{1916}, 65) \cdot \prod\nolimits_{act} \cdot \theta], \tag{3}$$

under the assumption the wife would claim on the husband's earnings. Then the cohort simulated Social Security income for a family with a male head claiming at age 65, $Z^{\overline{65}}$, was formed as a weighted average of (2) and (3), with the weights equal to share of unmarried and married male heads in each age-by-calendar year cell, respectively.

The solid line in Figure 11 shows $Z^{\overline{65}}$ by year of birth. The short-dashed line in the figure is actual Social Security income. Its trajectory tracks $Z^{\overline{65}}$ very closely from 1885-1896 and 1905-1916, indicating that much of the upward trend in family Social Security income across cohorts is due to general benefit increases. In (1), $Z^{\overline{65}}$ enters flexibly as a quartic function, so that the estimate of β will be identified from the cross-cohort variation in the early entitlement age, independent of benefit generosity. This is illustrated in the figure by divergence in the series for birth years 1897-1900, when early claiming is introduced.

VI. Estimation Results

Table 2 shows grouped Ordinary Least Squares (OLS) estimates of β, where the weights are based on the age-by-calendar-year cell sizes. There are 643 cells in the estimation sample, which ranges from ages 66-90.[5] Each column shows selected estimates from a separate regression. Row 1 of column 1 of Table 2 shows the estimate of β in (1) excluding the vector $\mathbf{X}$ of other controls, when the dependent variable, mean family Social Security Income, is measured in dollars. Controlling for age effects, calendar year effects, and age-65 benefit generosity, $\hat{\beta} = 71$, which says that a decrease of one year in the early entitlement age lowered Social Security income by 71 dollars. Based on the standard errors, clustered by year of birth and shown in parentheses, the null hypothesis of $\beta = 0$ can be rejected in favor of the alternative $\beta > 0$ at conventional significance levels. Relative to mean family Social Security income of $6,631 (shown in the second row of the table) for men born prior to 1897, and hence unaffected by, the 1961 Amendments, this estimate implies that a one-year reduction in the EEA lowered income by 1.1%, shown in the third row of the table.

Table 3 shows the fraction of men claiming at different ages from the 1916 birth cohort (the youngest in our sample), which, for the purposes of our analysis, we use to measure the long-run steady-state claiming distribution that resulted from the 1961 Amendments.[6] These data are

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[5] For ages greater than 90, the cells in the CPS became too thin, and, hence, were excluded from the estimation sample. Selected descriptive statistics for the sample are shown in Table 1.

[6] Ideally, we would use the claiming-age distribution for a birth year closer to the 1900 cohort, the pivotal cohort affected by the decrease in the EEA from the 1961 Amendments. Unfortunately, the 1900 cohort, and those that come shortly after, turn 65 in the years (1965 and later) in which individuals also are responding to the changes brought about by the 1965 Amendments. By choosing the 1916 cohort to measure the long-run steady-state claiming-age

compiled from claiming data published in various issues of the *Annual Statistical Supplement*. Just over 57% of men in the 1916 cohort took up early benefits, with 10.3% claiming at age 64, 15.3% at age 63, and 31.5% at age 62. The weighted-average reduction in claiming age below 65 for the 1916 cohort was therefore 1.4 years. Using this as a measure of the long-term reduction in claiming age implies that early claiming reduced Social Security income by 1.5%, as shown in the fourth row of Table 2.

Column 2 shows the same estimates with the vector $\mathbf{X}$ of other controls. The estimated impact of early claiming is larger: an increase of one year in the early entitlement age raised Social Security income by 107 dollars, which is larger, but not significantly different than, the results without controls. Relative to mean family Social Security income of $6,631, this estimate implies that a one-year reduction in the EEA lowered income by 1.6%, and that the average actual reduction in the early claiming age reduced Social Security income by 2.2%.

To examine how changes in the EEA affected the distribution of Social Security income, columns 3-7 of the table repeat the richest specification from column 2, but with the dependent variable, S, measuring selected percentiles of the Social Security income distribution. The estimates suggest the bulk of the impact of changes in the EEA were on the lower to middle part of the distribution, with small (and statistically insignificant) impacts in the tails. A one-year reduction in the EEA was associated with a reduction in Social Security income by 2.9% at the 25th percentile of the Social Security income distribution, and 2.1% at the median.

Figure 12 expands the estimates in Table 2 to the 1st-90th percentiles of the Social Security income distribution to illustrate more broadly the heterogeneity in the estimated impacts of

distribution, we are assured that those changes have played out. Our results are very similar if we use a cohort slight older than the 1916 cohort, e.g., the 1910 cohort.

reducing the EEA by one year. Specifically, the solid line in the figure shows the estimates of $(-)\beta$ in (1) when the dependent variable is the *q*th percentile value of the Social Security income distribution for age *a* in year *t*, where $q = 1,...,99$. The dashed lines demarcate the boundaries of the 95% confidence interval around the estimates. To keep the scale in the figure manageable, we do not show the estimates for the 91st-99th percentiles, which have very large confidence intervals.[7] Most of the statistically significant impacts of a one-year decrease in the EEA on Social Security income occur in the 20th-75th percentiles of the distribution, ranging from a $100-$150 decrease in income.

Table 4 shows a parallel set of reduced-form estimates as those in Table 2, but for total family income. At the mean, a one-year reduction in the EEA is associated with a reduction of total income of 62 dollars, which is not significant. There is a highly significant effect on income at the 25th percentile of $168, which is 2.3% of income at that point in the distribution. Figure 13 expands the estimates in Table 4 to the 1st-90th percentiles of the total income distribution to illustrate more broadly the heterogeneity in the estimated impacts of reducing the EEA by one year. Specifically, the solid line in the figure shows the estimates of $(-)\beta$ in (1) when the dependent variable is the *q*th percentile value of the total income distribution for age *a* in year *t*, where $q = 1,...,99$. Most of the statistically significant impacts of a one-year decrease in the EEA on total income occur in the bottom half of the distribution. The effect then becomes negative and insignificant at higher quantiles. Taken together, these results indicate that in the lower half of the income distribution there is no "crowd out" of Social Security income by other sources of income – so that the reductions in Social Security are translated to total income. However, at the upper

[7] They are available upon request.

end of the income distribution, we do not see such evidence – albeit with sufficient imprecision that we cannot rule out the same results as for Social Security income.[8]

To estimate the reduced-form impacts on poverty and inequality, Table 5 shows the reduced-form grouped OLS estimates of ρ from a specification isomorphic to (1):

$$P_{at} = \eta + \rho EEA_{at} + \boldsymbol{\xi}_a + \boldsymbol{\psi}_t + \boldsymbol{\delta}'\mathbf{X}_{at} + f(Z_{at}^{\overline{65}}) + \varepsilon_{at}. \tag{4}$$

In column 1, the dependent variable, P, is a measure of absolute poverty: the fraction of families below the federal poverty line. The focal parameter estimate in row 1, $\hat{\rho} = -0.0071$, indicates that a one-year decrease in the EEA raised the poverty rate about seven-tenths of a percentage point. Relative to an average poverty rate of 27.8 percentage points for cohorts prior to early claiming (shown in row 2), this estimate represents a 2.5% reduction in the poverty rate (i.e., $-0.0071/0.278 = -0.025$) and a 3.4% reduction based on the long-run actual reduction in the early claiming age (row 4).

In column 2, the dependent variable is relative poverty, measured as the fraction of male-headed elderly families with total family income below 40% of the median income of non-elderly families. The estimate suggests that a one-year reduction in the EEA was associated with an increase in relative poverty of around 0.7 percentage point, which is marginally significant. Relative to the mean relative poverty rate of 32.9 percentage points (row 2) for men born in cohorts ineligible for early claiming, this represents a 2.2% increase in the relative poverty rate (row 3) and a 2.9% increase based on the long-run actual reduction in the early claiming age (row 4).

[8] The quotient of the reduced-form total income estimate at the mean in column 1 of Table 4 to the Social Security income estimate at the mean in column 2 of Table 2 yields the IV estimate at the mean of 0.58 (with a standard error of 1.58). This implies that there is 42% crowd-out of Social Security income at the mean, but this estimate is not statistically different than either 0 or 1.

Column 3 shows the estimate of the reduced-form effect of eligibility on income inequality, measured as the 90-10 coefficient of variation in total family income. The estimate indicates that a one-year reduction in the EEA was associated with an increase in the coefficient of variation by 1.6% (row 3) and 2.2% (row 4), based on the long-run actual reduction in the early claiming age. These effects are marginally statistically different from zero, and suggest that reductions in elderly income from early claiming increased income inequality among the male-headed elderly families.

Finally, the first-stage estimates from (1) can be combined with the reduced-form estimates from (4) to yield instrumental variable (IV) estimates of the impact of Social Security income on poverty:

$$P_{at} = \mu + \phi S_{at} + \boldsymbol{\xi}_a + \boldsymbol{\psi}_t + \boldsymbol{\delta}'\mathbf{X}_{at} + f(Z_{at}^{\overline{65}}) + \upsilon_{at}, \tag{5}$$

where EEA is excluded from (5) as the instrument. Column 1 of Table 6 shows the IV estimate of ϕ for absolute poverty. The estimate is $\hat{\phi} = -0.066$ and indicates that a \$1000 increase in Social Security income is associated with a 6.7 percentage-point reduction in the poverty rate. Relative to mean Social Security income, this can be interpreted as an elasticity of the absolute poverty rate with respect to Social Security income of -1.6, shown in the second row. Columns 2 and 3 show similarly large elasticity estimates for relative poverty and income inequality, respectively.

VII. Summary and Caveats

This paper estimated the impact of the early entitlement age on later-life financial well-being by tracing the income and poverty status of different birth cohorts of men who had access to different potential claiming ages due to the Social Security Amendments of 1961. Reductions in the early entitlement age in the long-run lowered the average claiming age by 1.4 years, which lowered Social Security income for male-headed families in retirement by 1.5% at the mean, 3% at the median, and 4% at the 25th percentile of the Social Security income distribution. Early claiming also was associated with a decrease in total income at the bottom of the income distribution, and an increase in the elderly poverty and income inequality. Finally, for the 1885-1916 cohorts, the implied elasticity of poverty with respect to Social Security income for male-headed families is -1.6, similar to that estimated for male- and female-headed families pooled over the 1885-1933 cohorts in Engelhardt and Gruber (2005, 2006), using variation in benefit levels based on the Social Security notch.

These findings are tempered by a number of caveats. First, early claiming was introduced for men in a period of rapidly rising benefit generosity at all ages. These rising real benefits may have substantially offset the specific declines in retirement income from early claiming. That is, early claiming might have had larger impacts on poverty in the absence of these general benefit increases. This is important since future changes to entitlement ages will be done in a very different environment governing the trajectory of benefits.

Second, the introduction of early claiming for men in 1961 happened a long time ago; labor markets and retirement behavior for older workers have changed dramatically since then. In addition, as emphasized by Finkelstein and McKnight (2008), the cohorts under study here also experienced a large decline in old-age mortality rates, so that health was improving as these men aged. Any prospective changes in the EEA are unlikely to share these features of the broader economic and health environment facing older Americans.

Third, one avenue we did not explore directly was the extent to which early claiming might have been correlated with later-life mortality. In particular, there is a large literature in economics, demography, and sociology that documents an inverse relationship between income and mortality, which might suggest that the reduction in income for cohorts that claimed early might have found its way into higher later-life mortality for these cohorts. Two well-cited existing studies suggest this linkage is likely not important in our context. Finkelstein and McKnight (2008) documented a secular decline in mortality rates that was roughly parallel for younger vs. older elderly from the 1950s through the 1970s, with little evidence of differential mortality based on age across this time period, implying little cohort-specific changes in mortality for the cohorts in our study. This is consistent with the more direct evidence from Snyder and Evans (2006), who examined the impact of income on mortality by examining the income reductions from the Social Security notch. They found that the income reduction notch led to a reduction, not an increase in mortality.

Fourth, we examined whether there might be important heterogeneity in response for different subgroups of the elderly (by marital status, education, etc.), but were unable to draw firm conclusions given, in some cases, modest sample sizes. This is an important direction for further analysis.

Finally, although income poverty is an important policy metric, it is not the only measure of well-being. In addition to health and changes in material well-being, such as non-durable consumption and housing, men who claimed benefits early also consumed more leisure than those who claimed later, compensating them to at least some extent for the loss in income from the actuarial reduction. A particularly important question that remains is whether early claiming decisions are time-inconsistent. That is, do men (and their family members) who claimed early and ended up in retirement and old age with less income and more impoverished regret the decision to have claimed early? Analysis of these types of issues is important for a richer understanding of the welfare implications of changes to Social Security claiming.

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Table 1

Sample Means

Variable	(1) Mean	(2) Standard Deviation
Family Social Security Income Per Equivalent	7,624	1,464
Median Family Total Income Per Equivalent	17,627	3,190
Absolute Poverty Rate	0.194	0.085
Relative Poverty Rate	0.259	0.080
90-10 Coefficient of Variation in Total Family Income	1.536	0.211
Early Claiming Age	62.9	1.3
Simulated Social Security Income for Age-65 Claiming	10,386	2,548
Veteran	0.727	0.126
White	0.913	0.029
High School Graduate	0.182	0.069
Some College	0.079	0.037
College Graduate	0.055	0.024
More than College	0.044	0.022
Married	0.689	0.113
Divorced or Separated	0.034	0.017
Widowed	0.232	0.123
Medicare Eligible	0.628	0.484
Disability Insurance Eligible	0.844	0.363
Quarters of Coverage for Age-72 Benefits	16.2	16.8

Notes: The table shows means and standard deviations for selected variables from the CPS data set described in text.

Table 2

OLS Estimation Results for Family Social Security Income; Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Dependent Variable: Social Security Income						
	Mean	Mean	10^{th} Percentile	25^{th} Percentile	Median	75^{th} Percentile	90^{th} Percentile
Parameter Estimates							
Early Entitlement Age	71	107	26	134	144	87	81
	(32)	(31)	(100)	(35)	(32)	(40)	(69)
Estimated Impacts							
Dependent Variable Mean Prior to Availability of Early Claiming	6,631	6,631	2,471	4,700	6,900	9,833	11,817
Estimated Impact of One-Year Reduction in Early Entitlement Age as a % of Dependent Variable Mean Prior to Availability of Early Claiming	-1.1	-1.6	-1.1	-2.9	-2.1	-0.9	-0.7
Estimated Impact of Weighted-Average Actual Reduction in Claiming Age for 1916 Cohort (1.35 Years Earlier) as a % of Dependent Variable Prior to Availability of Early Claiming	-1.5	-2.2	-1.4	-3.9	-2.8	-1.2	-0.9
Additional Controls							
Demographics and Veteran Status	No	Yes	Yes	Yes	Yes	Yes	Yes
Medicare and DI Effects	No	Yes	Yes	Yes	Yes	Yes	Yes
Quarters of Coverage for Age-72 Benefits	No	Yes	Yes	Yes	Yes	Yes	Yes
Age and Calendar-Year Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Quartic Function of Simulated Social Security Income for Age-65 Claiming	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: N = 643. The first row of table shows the parameter estimate of beta in (1), which is the impact of a one-year increase in the early entitlement age on the Social Security income measure. Each column represents a different regression, based on a different measure of Social Security income. All regressions include the full set of age and year dummies, and a quartic in simulated Social Security income for age 65 claiming to control for trends in benefit generosity across cohorts that were not due to early claiming. Columns 2-7 add in additional controls for the percentage in age/year cell that are: veteran status; white; high school graduate; some college; college graduate; advanced degree; plus eligibility for Medicare and DI, and quarters of coverage for age-72 benefits, respectively. Standard errors clustered by year of birth are shown in parentheses. The second row shows the dependent variable mean for cohorts not eligible for early claiming. The third row is (minus) the quotient of rows 1 and 2, and represents the estimated impact of a one-year reduction in the claiming age as a % of the dependent variable for the cohorts not exposed to early claiming. The fourth row is (minus) the quotient of rows 1 and 2, multiplied by 1.4, and represents the estimated impact of the long-run actual reduction in the claiming age (as measured by the 1916 cohort) as a % of the dependent variable for the cohorts not exposed to early claiming.

Table 3

Fraction of Men Born in 1916 Claiming Social Security Retired Worker Benefits by Age

Age	Fraction Claiming
62	0.315
63	0.153
64	0.103
65	0.355
66	0.044
67	0.011
68	0.006
69 or older	0.012

Note: Authors' calculation from various issues of the *Annual Statistical Supplement*.

Table 4

OLS Estimation Results for Total Family Income; Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent Variable: Total Income					
	Mean	10th Percentile	25th Percentile	Median	75th Percentile	90th Percentile
Early Entitlement Age	62	52	168	82	-51	-212
	(175)	(50)	(77)	(81)	(192)	(444)
Dependent Variable Mean Prior to Availability of Early Claiming	14,295	5,228	7,344	10,355	20,523	33,565
Estimated Impact of One-Year Reduction in the Early Entitlement Age as a % of Dependent Variable Mean Prior to Availability of Early Claiming	-0.4	-1.0	-2.3	-0.8	0.2	0.6
Estimated Impact of Weighted-Average Actual Reduction in Claiming Age for 1916 Cohort (1.35 Years Earlier) as a % of Dependent Variable Prior to Availability of Early Claiming	-0.6	-1.3	-3.1	-1.1	0.3	0.8
Additional Controls						
Demographics and Veteran Status	Yes	Yes	Yes	Yes	Yes	Yes
Medicare and DI Effects	Yes	Yes	Yes	Yes	Yes	Yes
Quarters of Coverage for Age-72 Benefits	Yes	Yes	Yes	Yes	Yes	Yes
Age and Calendar-Year Effects	Yes	Yes	Yes	Yes	Yes	Yes
Quartic Function of Social Security Income for Age-65 Claiming	Yes	Yes	Yes	Yes	Yes	Yes

Notes: N = 643. The first row of table shows the parameter estimate of beta in (1), which is the impact of a one-year increase in the early entitlement age on the total income measure. Each column represents a different regression, based on a different measure of total income. All regressions include the full set of age and year dummies, a quartic in simulated Social Security income for age 65 claiming to control for trends in benefit generosity across cohorts that were not due to early claiming, and additional controls for the percentage in age/year cell that are: veteran status; white; high school graduate; some college; college graduate; advanced degree; plus eligibility for Medicare and DI, and quarters of coverage for age-72 benefits, respectively. Standard errors clustered by year of birth are shown in parentheses. The second row shows the dependent variable mean for cohorts not eligible for early claiming. The third row is (minus) the quotient of rows 1 and 2, and represents the estimated impact of a one-year reduction in the claiming age as a % of the dependent variable for the cohorts not exposed to early claiming. The fourth row is (minus) the quotient of rows 1 and 2, multiplied by 1.4, and represents the estimated impact of the long-run actual reduction in the claiming age (as measured by the 1916 cohort) as a % of the dependent variable for the cohorts not exposed to early claiming.

Table 5

OLS Estimation Results for Absolute Poverty, Relative Poverty, and Income Inequality, Respectively; Standard Errors in Parentheses

	(1)	(2)	(3)
		Dependent Variable:	
	Absolute Poverty: Fraction of Families Below Federal Poverty Line	Relative Poverty: Fraction of Families with Total Income Below 40% of Non-Elderly Median Income	Income Inequality: 90-10 Coefficient of Variation for Total Family Income
Parameter Estimates			
Early Entitlement Age	-0.0071	-0.0072	-0.024
	(0.0035)	(0.0045)	(0.015)
Estimated Impacts			
Dependent Variable Mean Prior to Availability of Early Claiming	0.278	0.329	1.486
Estimated Impact of One-Year Reduction in the Early Entitlement Age as a % of Dependent Variable Mean Prior to Availability of Early Claiming	2.5	2.2	1.6
Estimated Impact of Weighted-Average Actual Reduction in Claiming Age for 1916 Cohort (1.35 Years Earlier) as a % of Dependent Variable Prior to Availability of Early Claiming	3.4	2.9	2.2
Additional Controls			
Demographics and Veteran Status	Yes	Yes	Yes
Medicare and DI Effects	Yes	Yes	Yes
Quarters of Coverage for Fully Insured and Age 72 Benefits	Yes	Yes	Yes
Age and Calendar Year Effects	Yes	Yes	Yes
Quartic Function of Simulated Social Security Income for Age-65 Claiming	Yes	Yes	Yes

Notes: N = 643. The first row of table shows the parameter estimate of phi in (4), which is the impact of a one-year increase in the early entitlement age on the poverty or inequality measure. Each column represents a different regression, based on a different measure of poverty or inequality. All regressions include the full set of age and year dummies, a quartic in simulated Social Security income for age 65 claiming to control for trends in benefit generosity across cohorts that were not due to early claiming, and additional controls for the percentage in age/year cell that are: veteran status; white; high school graduate; some college; college graduate; advanced degree; plus eligibility for Medicare and DI, and quarters of coverage for age-72 benefits, respectively. Standard errors clustered by year of birth are shown in parentheses. The second row shows the dependent variable mean for cohorts not eligible for early claiming. The third row is (minus) the quotient of rows 1 and 2, and represents the estimated impact of a one-year reduction in the claiming age as a % of the dependent variable for the cohorts not exposed to early claiming. The fourth row is (minus) the quotient of rows 1 and 2, multiplied by 1.4, and represents the estimated impact of the long-run actual reduction in the claiming age (as measured by the 1916 cohort) as a % of the dependent variable for the cohorts not exposed to early claiming.

Table 6

IV Estimation Results for Absolute Poverty, Relative Poverty, and Income Inequality, Respectively; Standard Errors in Parentheses

	(1)	(2)	(3)
	Dependent Variable:		
	Absolute Poverty: Fraction of Families Below Federal Poverty Line	Relative Poverty: Fraction of Families with Total Income Below 40% of Non-Elderly Median Income	Income Inequality: 90-10 Coefficient of Variation for Total Family Income
Parameter Estimates			
Mean Social Security Income	-0.066	-0.067	-0.227
	(0.0024)	(0.033)	(0.167)
Estimated Elasticities			
Outcome Elasticity with respect to Mean Social Security Income	-1.6	-1.3	-1.0
Additional Controls			
Demographics and Veteran Status	Yes	Yes	Yes
Medicare and DI Effects	Yes	Yes	Yes
Quarters of Coverage for Fully Insured and Age 72 Benefits	Yes	Yes	Yes
Age and Calendar Year Effects	Yes	Yes	Yes
Quartic Function of Simulated Social Security Income for Age-65 Claiming	Yes	Yes	Yes

Notes: N = 643. The first row of table shows the IV parameter estimate of rho in (5), which is the impact of a $1000 increase in Social Security income on the poverty or inequality measure, where the cohort's early entitlement age is used as the instrument. Each column represents a different regression, based on a different measure of poverty or inequality. All regressions include the full set of age and year dummies, a quartic in simulated Social Security income for age 65 claiming to control for trends in benefit generosity across cohorts that were not due to early claiming, and additional controls for the percentage in age/year cell that are: veteran status; white; high school graduate; some college; college graduate; advanced degree; plus eligibility for Medicare and DI, and quarters of coverage for age-72 benefits, respectively. Standard errors clustered by year of birth are shown in parentheses. The second row shows the implied elasticity of the poverty or inequality measure with respect to Social Security income.

Figure 1. Number of Retired-Worker Awards for
Both Sexes Pooled

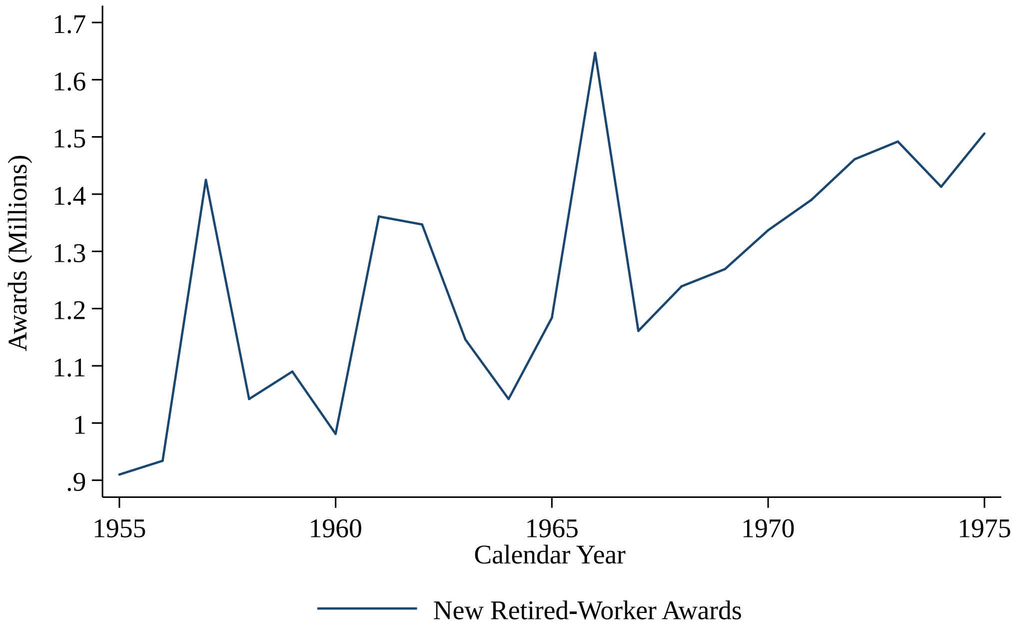

Figure 2. Number of New Retired-Worker Awards to Women
by Age Group

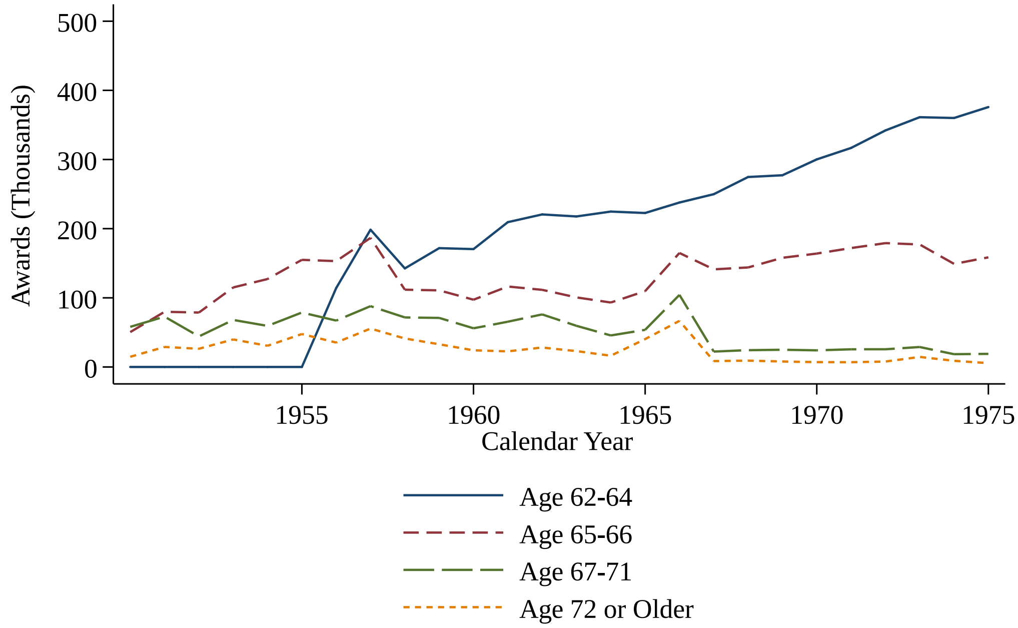

Figure 3. Number of New Retired-Worker Awards to Men
by Age Group

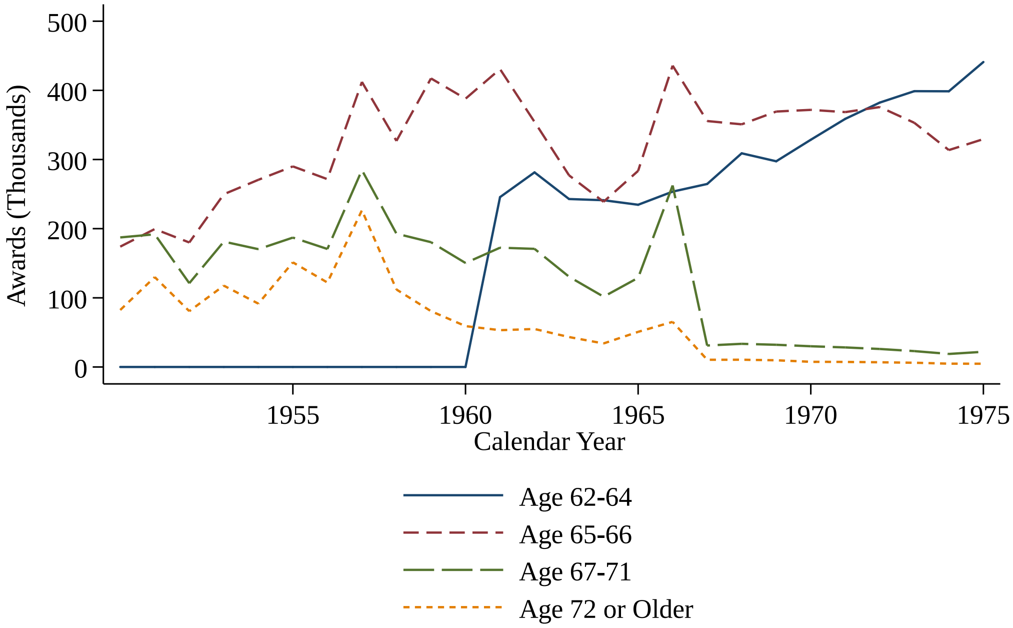

Figure 4. Average Annual Amount of New Retired-Worker Award and Percent of New Awards from Early Claiming for Men

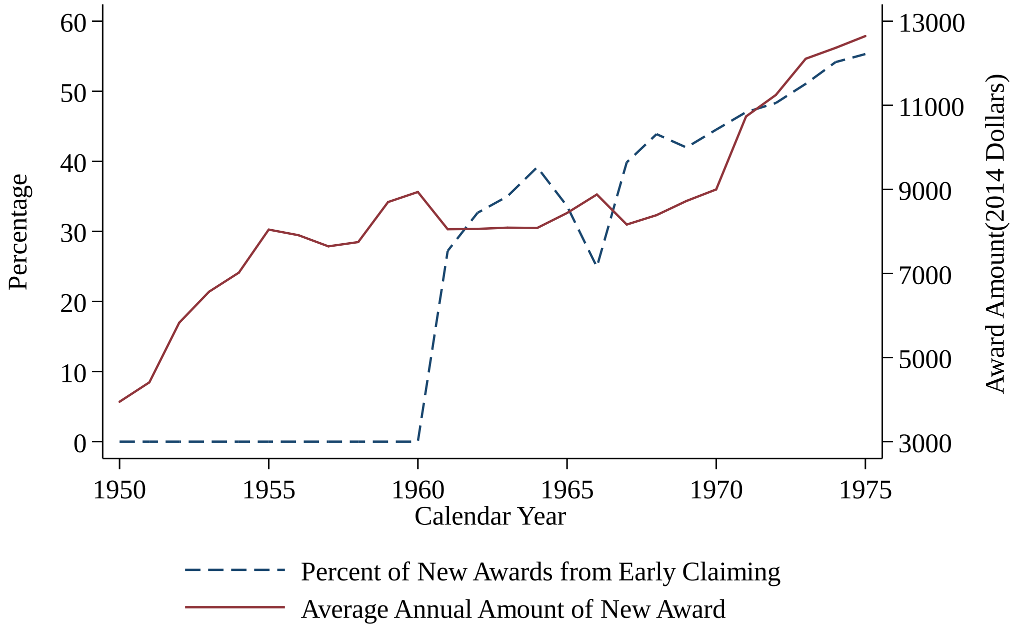

Figure 5. Fraction of New Retired-Worker Awards to Men
by Claiming Age and Year of Birth

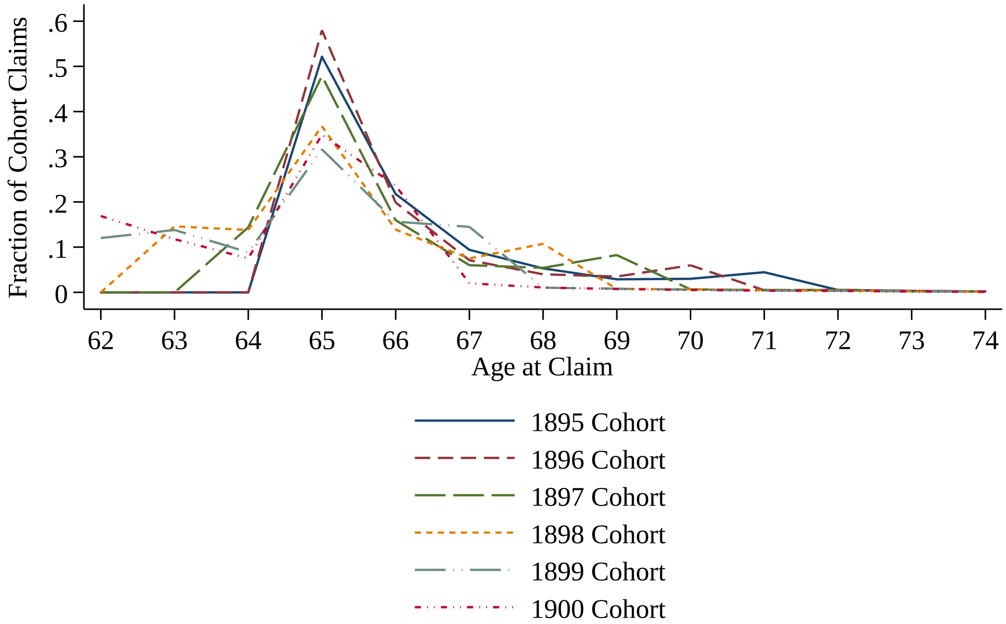

Figure 6. Social Security and Total Income for Male-Headed Elderly Families by Year of Birth

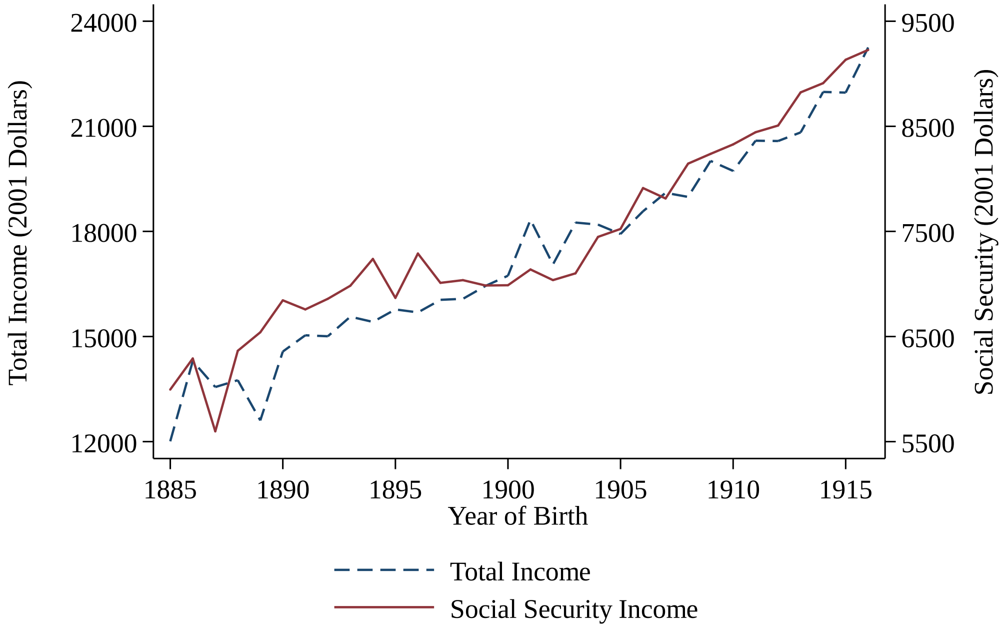

Figure 7. Social Security Income and Absolute Poverty of Male-Headed Elderly Families by Year of Birth

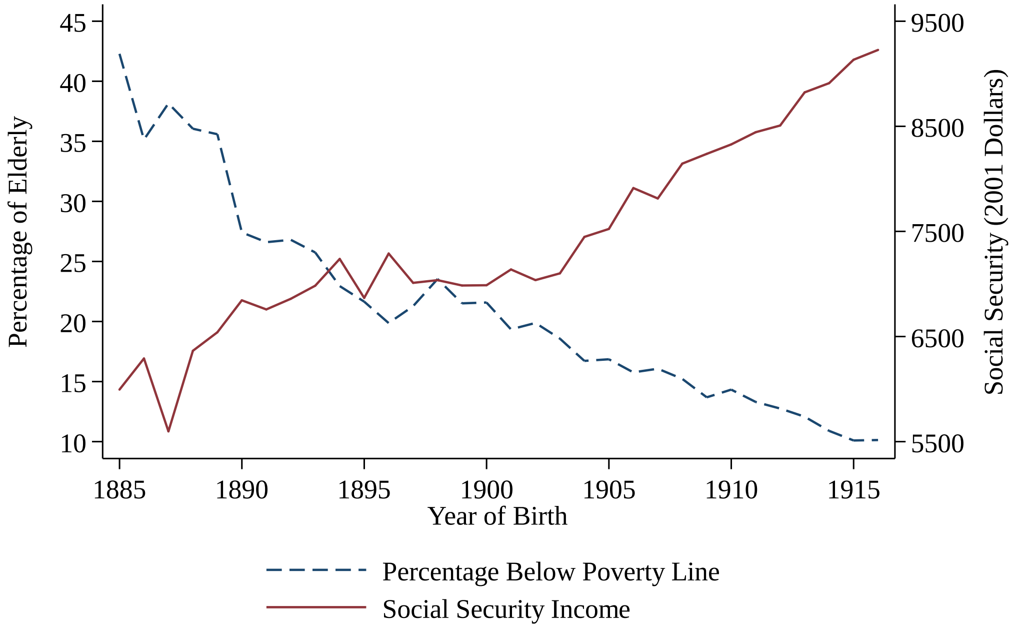

Figure 8. Social Security Income and Relative Poverty of Male-Headed Elderly Families by Year of Birth

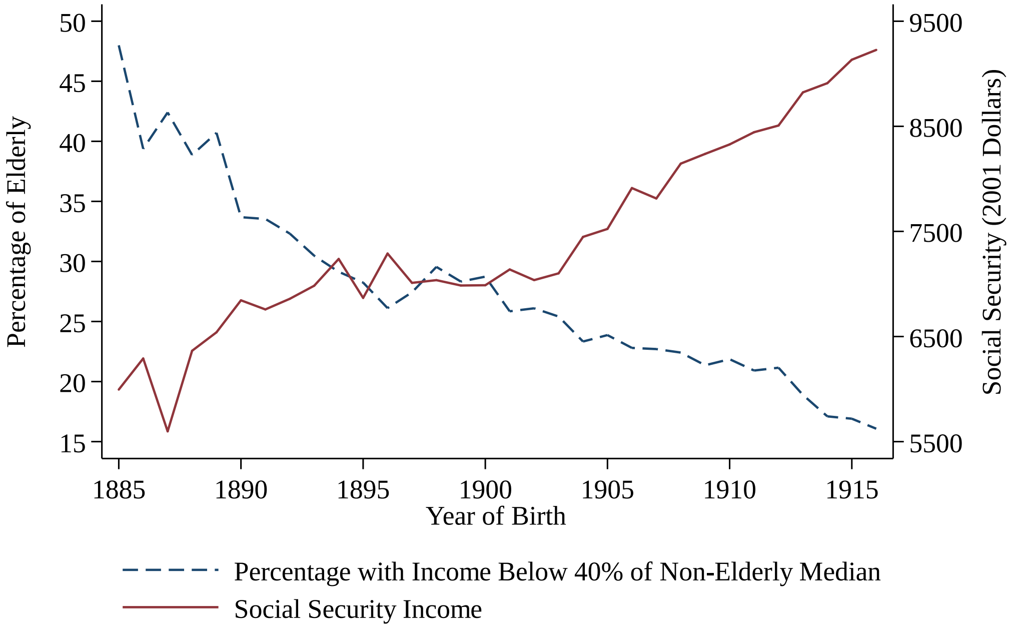

Figure 9. Social Security and Income Inequality for Male-Headed Elderly Families by Year of Birth

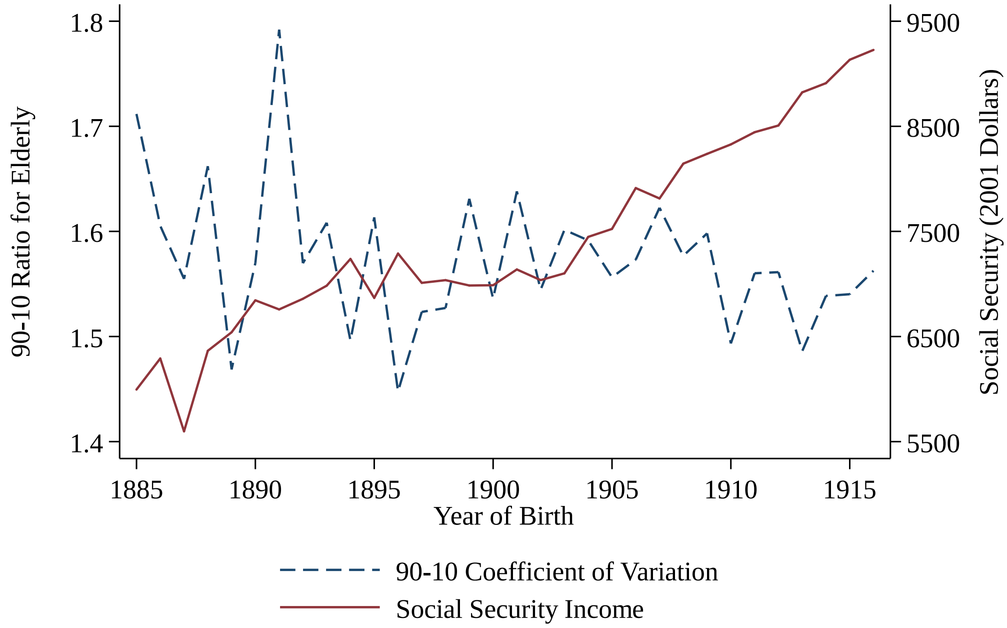

Figure 10. Primary Insurance Amount at Different Claiming Ages
by Year of Birth for 1916 Cohort Median Male Earnings

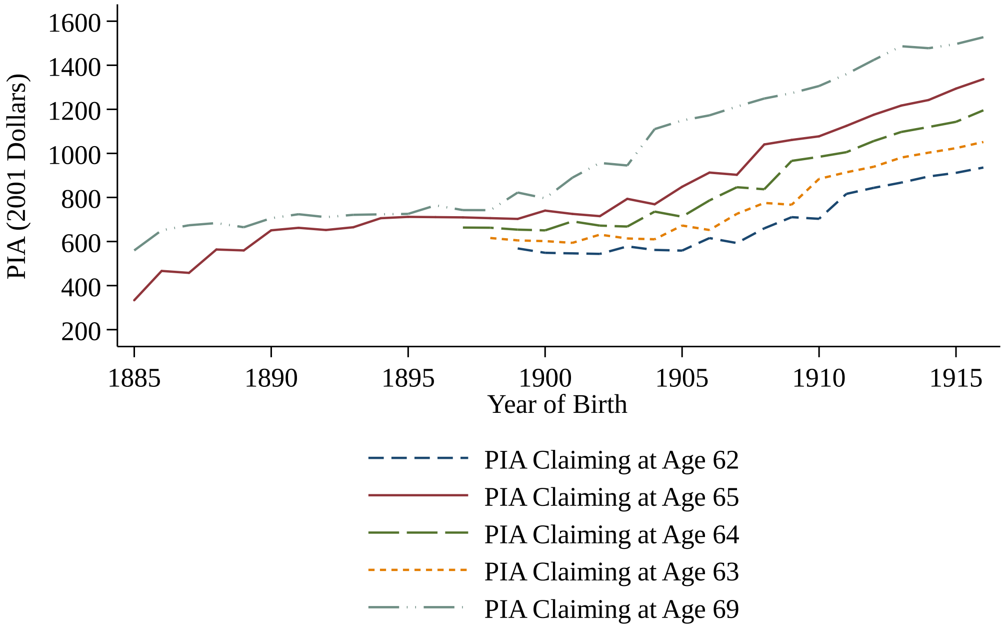

Figure 11. Simulated and Actual Social Security Income for Male-Headed Elderly Families by Year of Birth

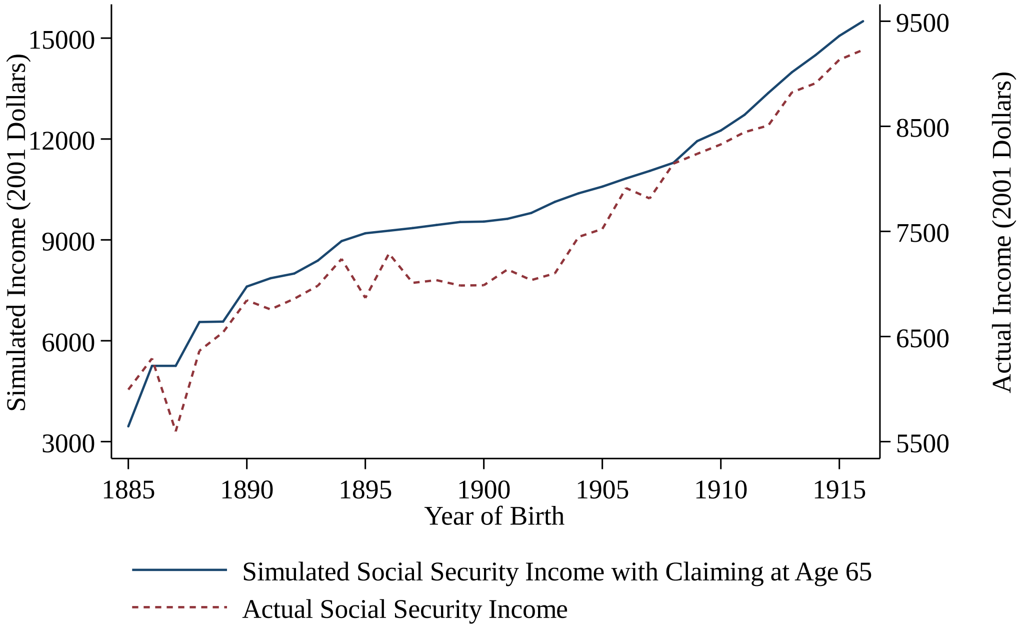

Figure 12. Impact of Reducing the Early Entitlement Age by One Year on Social Security Income for Percentiles of the Social Security Income Distribution

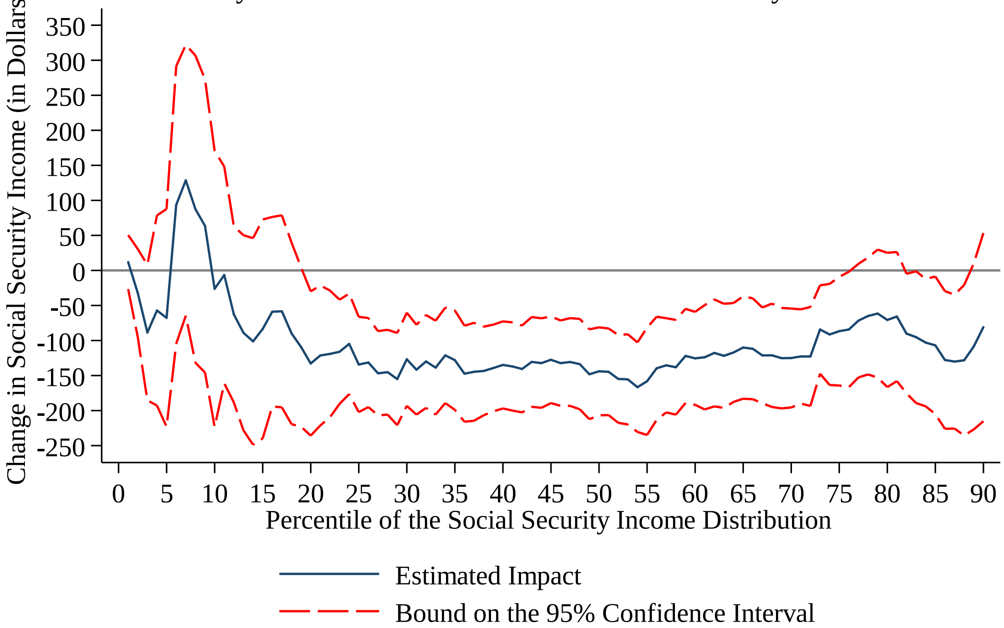

Figure 13. Impact of Reducing the Early Entitlement Age by One Year on Total Income for Percentiles of the Total Income Distribution