The Introduction of Social Security and Elderly Mortality*

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Abstract

The introduction of Social Security in the United States in 1935 and its subsequent expansion were associated with transformative changes in the lives of the elderly, such as falling male labor force participation at older ages and rising independent living among older women and men. In this paper, we investigate how the early Social Security program affected longevity, a fundamental dimension of well-being and an important determinant of the fiscal sustainability of old-age support programs. We combine an empirical approach that exploits within-occupation, cross-industry differences in Social Security coverage in the early years of the program with a dataset connecting men and women in the 1930 full-count Census to their death records. We find that men who were more likely to have received Social Security, or received it earlier, lived longer as a result. In contrast, we find no evidence that the early Social Security program affected the longevity of their wives or that of unmarried women. Evidence based on the task content of occupations from the 1965 *Dictionary of Occupational Titles* suggests that earlier exit from jobs that were physically demanding or involved poor working conditions was an important channel through which Social Security extended the lives of older men in the mid-20th century.

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1 Introduction

The introduction of Social Security in 1935 and its subsequent expansion were associated with transformative shifts in the lives of the elderly, including large declines in labor force participation at older ages and rising rates of independent living among older men and women (Costa, 1998; Fetter and Lockwood, 2018). In this paper, we investigate how the early Social Security program affected the longevity of its recipients, a fundamental dimension of well-being and an important determinant of the fiscal cost of old-age support programs.

Whether Social Security increased or decreased longevity is ambiguous a priori. On the one hand, Social Security provided a significant new source of income to elderly recipients, which could have improved health through better nutrition, housing, and medical care. On the other hand, by inducing earlier exit from the labor force, Social Security may have reduced the social connections, sense of purpose, and structure that work provides—factors that some evidence suggests are protective of health (Fitzpatrick and Moore, 2018; Snyder and Evans, 2006).

We use an empirical approach for estimating the causal impact of Social Security introduced by Fetter, Lockwood and Mohnen (2026). This approach makes use of the fact that employment in some industries was covered when the program was introduced in 1935, while similar employment in other industries was covered only in 1950 or later as the program was expanded. For example, work as a janitor in a private, for-profit firm was covered in 1935, while work as a janitor in a nonprofit hospital or educational institution was not covered until 1950. As a result, due to their industry of employment, individuals with similar jobs sometimes differed in the likelihood of ever becoming eligible for Social Security, or became eligible for Social Security benefits at different ages. Moreover, differences in coverage year had differential "bite" across cohorts since eligibility was conditional on being 65 or older; among workers staying in 1950-covered employment, one who turned 65 in the early 1950s was much less affected than one who turned 65 ten years earlier. We leverage this within-occupation, cross-industry variation for different cohorts to compare the longevity of otherwise-similar individuals who differed in the timing and likelihood of their Social Security eligibility.

We combine this empirical approach with a dataset that connects men and women in the 1930 full-count Census to death information available on the Family Tree at FamilySearch.org. Most of this death information is from death records on FamilySearch, but it also comes from records on other genealogical websites or the personal knowledge of the contributor. The 1930 Census allows us to predict the age at which a worker (and their spouse) would likely become eligible for Social Security, based on their year of birth and their employment information prior to the passage of the Social Security Act. Information from death records allows us to calculate each individual's age at death, and come from a variety of sources, including federal and state death

records, obituaries, cemetery listings, and funeral home records. To investigate the role of work as a channel through which Social Security affected longevity, we additionally draw on information from the 1965 *Dictionary of Occupational Titles* on the physical demands and hazards associated with different occupations.

We find that men who were more likely to have received Social Security, or received it earlier, lived longer as a result. Our preferred estimate implies that a ten-year difference in Social Security exposure—roughly the difference between 1935-covered and 1950-covered workers from the 1877 birth cohort—increased men's age at death by approximately 0.45 years (about 5.4 months), conditional on surviving to age 55. The effect is roughly five times larger for Black men (about 1.6 years) than for white men (about 0.3 years). In contrast, we find no evidence that the early Social Security program affected the longevity of men's wives or the longevity of unmarried women eligible for Social Security based on their own employment.

Several pieces of evidence point to earlier exit from physically demanding or hazardous jobs as an important channel for the longevity gains we observe among men. First, men predicted to be eligible for Social Security earlier were significantly less likely to be in the labor force by 1950—a notable finding in itself, given the paucity of evidence directly linking Social Security to the rise in male retirement over the mid-20th century (see, e.g. Fetter and Lockwood, 2018). Second, both the longevity effects and the labor force participation effects of Social Security are concentrated among men in occupations that rank highly on an index of physical demands and hazardous working conditions constructed from the 1965 *Dictionary of Occupational Titles*, while the effects for men in other occupations are small and statistically insignificant. Third, that Social Security extended men's lives but not their wives' lives is consistent with a channel—the physical demands and hazards of work—whose health consequences are borne by the worker rather than by the worker's spouse. That physical demands or hazards of work was a key channel is also consistent with larger effects for Black men, among whom these types of jobs were more prevalent, and with modest or no effects for working women, among whom these types of jobs were less prevalent.

This paper most directly contributes to the literature on the relationship between Social Security and longevity, which has arrived at differing conclusions depending on the source of variation and the time period studied. Snyder and Evans (2006) exploit variation from the Social Security "Notch" and find that men with higher benefits had higher mortality rates, which they attribute to lower labor force participation and greater social isolation. However, Noghanibehambari and Fletcher (2025) study the same policy variation using death records with a longer follow-up period and find a positive longevity effect of higher benefits. Engelhardt et al. (2022) study the introduction of early (age-62) claiming for men in 1961 and find modest negative effects of early claiming on longevity overall, though they note that this policy change involved a tradeoff between earlier benefits and lower monthly amounts, and had heterogeneous effects by age, shifting mortality

earlier in some age ranges and later in others. Fitzpatrick and Moore (2018) document spikes in mortality at age 62 linked to exit from the labor force.

In addition to using a different source of variation and our ability to observe completed longevity, our analysis differs from this prior work in studying the initial introduction and expansion of the program, when the net transfers to recipients were far larger than those associated with later policy changes, and when the composition of employment among older men differed from more recent periods. Our finding that the longevity effects of Social Security are concentrated among workers in physically demanding and hazardous occupations has implications for understanding how the effects of Social Security on longevity may have changed over time. As the share of workers in such occupations has declined over the course of the twentieth century, the marginal worker induced to retire by Social Security today is less likely to be in a job that is directly harmful to health. This may help to reconcile our finding of a positive longevity effect of the early Social Security program with more recent evidence suggesting that retirement can be harmful to health in some contexts. It also resonates with Costa (2000), who attributes an important portion of the twentieth-century decline in chronic disease among men to the shift from manual to white-collar jobs.

The contrast between our findings for men and women speaks to the literature on the divergence between male and female life expectancy over the mid-twentieth century. Our findings stand in notable contrast to overall trends during this period, when female life expectancy grew substantially while male life expectancy was relatively stagnant. Beltrán-Sánchez et al. (2015) attribute this divergence in large part to higher smoking rates among men, while Goldin and Lleras-Muney (2019) argue that it partly reflects the waning legacy of childhood infectious disease, which had disproportionately harmed women's health. Our findings, to some degree, heighten the contrast between men's and women's trends in life expectancy by suggesting that Social Security—by leading to greater longevity benefits for men—may have partially counteracted the forces that were otherwise restraining gains in male longevity.

2 Background on the early Social Security Program

Our approach to investigating the effect of Social Security on longevity of its early recipients follows Fetter, Lockwood and Mohnen (2026), and is based on the introduction of Social Security and its early expansions in types of employment covered under the program.[1] The Social Security Act was signed into law in 1935, significantly amended in 1939, and began issuing monthly retirement and survivors' benefits under the Old Age and Survivors Insurance (OASI) Program in 1940. We

[1] We discuss the key elements of eligibility and other aspects of the program here; further details can be found in Fetter, Lockwood and Mohnen (2026).

refer to these Old Age and Survivors Insurance (OASI) benefits as "Social Security" throughout. Under the 1939 amendments, eligibility for monthly retirement benefits depended on accumulating a sufficient number of quarters of work in covered employment, with each qualifying quarter requiring earnings above a minimum threshold.[2] Under the 1939 rules, for instance, a worker could qualify by earning enough in covered employment during at least half the quarters between 1936 (or age 21, if later) and age 65 or death, subject to a six-quarter minimum.

A feature of the program's early history that is central to our empirical strategy is that different types of employment became covered at different times. The 1935 Act and 1939 Amendments covered only wage and salary workers—excluding the self-employed—and even among wage and salary workers, several categories were initially excluded: domestic workers in private homes, agricultural laborers, employees of nonprofit organizations, and workers employed by state, local, and federal governments.[3] For employment covered in the 1935 Act, coverage began—and workers could begin accumulating quarters of coverage—on January 1, 1937.

The scope of covered employment expanded considerably beginning with the 1950 Amendments. Agricultural and domestic wage workers became subject to compulsory coverage provided they met regularity-of-employment thresholds. Employees of nonprofit organizations and state and local government workers who were not already enrolled in a retirement system became eligible for coverage on an elective basis. Critically, these extensions applied only to work performed from 1951 onward—coverage was not made retroactive. A worker employed exclusively in one of these newly covered sectors would therefore not begin accruing quarters of coverage until 1951, roughly 14 years after workers in originally covered industries had started doing so. Although participation was initially partially elective for nonprofit and government employees, adoption was rapid. As of 1951, 56% of employees of nonprofit organizations were covered under Social Security, rising to 76% by 1960 (Subcommittee on Social Security, 1976, p. 27). Adoption was slightly slower among employees of state and local governments, with about 11 percent of all state and local government employees covered by Social Security at the beginning of 1951, but by 1958, the share had risen to 54 percent (Kestenbaum, 1982).[4]

Workers who met the eligibility requirements could begin receiving monthly retirement benefits at age 65, though an earnings test withheld payments in months when earnings exceeded a

[2]The 1935 Act originally used a different eligibility formula, requiring both sufficient earnings and at least one day of covered employment in each of five calendar years after 1936 and before age 65. The 1939 Amendments replaced this with the quarters-of-coverage system.

[3]See Fetter, Lockwood and Mohnen (2026) for a comprehensive account of the original exclusions and the rationale behind them. Work performed at age 65 or older was not covered under the original act but became covered under the 1939 Amendments, effective January 1, 1939.

[4]Later amendments further expanded government coverage: the 1954 Amendments extended elective coverage to state and local employees already under a retirement system, and 1956 amendments further liberalized these provisions.

certain threshold. In the 1939 Amendments, this threshold was set at $15, which corresponded to roughly $230 in 2010 dollars in 1939, and to about $135 in 1950; the earnings test was liberalized in 1950 and subsequent amendments. Spouses of retired male workers received supplementary benefits equal to half the primary amount, while widows received survivor's benefits at 75 percent of the primary amount. Monthly benefit levels depended on cumulative earnings in covered employment; replacement rates for median earners were roughly 20 percent during the 1940s, rising to 25–30 percent through the 1950s and most of the 1960s (SSA Office of the Chief Actuary, 2013, Supplementary Table V.C7).

It has long been recognized that the earliest cohorts of beneficiaries received transfers that were extraordinarily generous relative to the taxes they had paid (e.g., Moffitt, 1984; Schieber and Shoven, 1999; Gruber, 2013). To illustrate the magnitude, consider a married couple in which the husband was born in 1880, worked continuously in originally covered employment at average earnings until age 65, and the wife (born in 1883) did not work outside the home. If both spouses survived to their gender-specific remaining life expectancy at age 65, the net present value at retirement of the couple's lifetime benefits minus taxes paid would have been approximately $84,000 in 2010 dollars, nearly double the 1940 median home value of about $45,774 in 2010 dollars.

3 Data

3.1 Linking the 1930 Census to death records, claims records, and the 1950 Census

Our main analysis relies on a dataset connecting individuals in the 1930 full-count Census (Ruggles et al., 2024) to death records from a genealogy website, FamilySearch.org. FamilySearch is a wiki-style online platform that helps users build family trees and gather information about their ancestors by linking individual "profiles" to a wide collection of historical records, including vital records (birth, marriage, death), census records, and a variety of other records. Thanks to FamilySearch's large and active user base (around 25 million average monthly site visits in 2025), the resulting dataset, called the "Family Tree," contains profiles for over 1.8 billion deceased individuals (FamilySearch, 2026). In this paper, we leverage an extract of profiles connected to the 1930 Census and that also include year of death information. The year of death can come from multiple sources, including federal and state death records, obituaries, cemetery listings, funeral home records, or the personal memory of the contributor. The Family Tree is one of the most comprehensive large-scale datasets available to study the determinants of longevity in this period, and has been used in other work (e.g., Lleras-Muney et al., 2025a,b; Goodman-Bacon et al., 2025a,b; Black et al., forthcoming).

Following Fetter, Lockwood and Mohnen (2026), most of our analysis focuses on men born between 1875 and 1888 who were covered by Social Security in 1935 or in 1950 based on their

employment information in 1930, excluding the self-employed.[5] After excluding individuals with implausible time of death information (people who reportedly lived more than 110 years or whose death is listed as occurring prior to 1930 even though they are linked to a person in the 1930 Census), we end up with just over 1.7 million men, which represents 35 percent of our population of interest in 1930. Our sample also includes year of death information for around 1.2 million spouses of men in our linked sample through the combination of within-household spousal links in the 1930 Census and links to death records in the Family Tree (representing 31.4 percent of spouses of men in our population of interest).

Appendix Figure A.1 shows that the distributions of birth cohorts in our linked sample are very similar to population analogs, for men, women, and within couples. The first two columns in Appendix Table A.1 show average characteristics in our population of interest in 1930 and the subset of men who appear in our linked sample. While our linked sample is broadly representative of the population, it is notably selected on race and birthplace: Black Americans and foreign-born individuals only make up 1 percent and 14.2 percent of our linked sample, whereas the corresponding population shares are 7.8 and 30.7 percent, respectively. In addition, higher-socioeconomic status individuals are slightly overrepresented in our linked sample, as measured by literacy status (97.6 vs. 94 percent), homeownership (54.9 vs. 47.2 percent), and white-collar employment (31.1 vs. 24.3 percent). These differences reflect some combination of the demographics of the FamilySearch user base, the imperfect coverage of death records, and the quality of the information that users have at their disposal to connect different records together. To address these imbalances, we use the procedure from Bailey et al. (2020) to construct inverse propensity score weights that re-weight men in our sample to be representative of our population of interest in terms of demographics (cohort, race, birthplace) and various economic characteristics (literacy, farm status, urban status, homeownership status, radio ownership status). By construction, the resulting weighted means in column 3 of Appendix Table A.1 are very close to the population means in column 1.

The top four panels in Appendix Figure A.2 show the distribution of age at death and year of death for men and women in our linked sample, while the bottom two panels show within-couple differences in age at death and year of death. We highlight a few key empirical facts. First, the average age at death in our sample is 72.8 years for men and 76.3 years for their wives, consistent with women living longer than men. The relatively high average age at death reflects that we are implicitly conditioning on men or husbands surviving to 1930 (i.e., surviving to age 42 to 55, depending on the cohort). As a benchmark, life expectancy at birth for men and women

[5]One slight departure from Fetter, Lockwood and Mohnen (2026) is that since our focus is on the recipients themselves as opposed to their children, we drop the restriction that men be ever-married in 1930. As a result, our sample includes single men and couples who never had children.

born in 1900 was 51.5 and 58.3 years, respectively, according to life tables from the Social Security Administration (Bell and Miller, 2005). Second, the gender gap in year of death is twice as large as the gender gap in age at death, since women tended to be a few years younger than their husbands in this period (by around 4.3 years in our sample). Third, despite that around two thirds of couple members are born within 5 years of each other, it is quite common for couple members to die many years apart (only a quarter die within 5 years of each other). This reflects the significant variance of longevity (Compton and Pollak, 2021). Lastly, panel (f) reveals excess mass at 0 in within-couple year of death differences, meaning that a disproportionate number of couple members die in the same year. This empirical regularity, often referred to as the "widowhood effect," has been documented in the literature (Elwert and Christakis, 2008).

We make use of two additional linked samples in our analysis. First, we leverage links between the 1930 Census and Numident claims records for 420,000 men in our population of interest, which are taken from Fetter, Lockwood and Mohnen (2026). This allows us to estimate the relationship between the timing of predicted eligibility for Social Security and subsequent claiming behavior (both the probability of ever claiming and the timing of claiming, conditional on ever claiming).[6] Second, we leverage links between the 1930 and 1950 Censuses for 1.8 million men in our population of interest, which come from the IPUMS Longitudinal Multigenerational Panel (Ruggles et al., 2025a). This allows us to observe the labor force participation status of early recipients 10 years after the first monthly Social Security benefits were issued, and test whether earlier predicted eligibility is associated with earlier exit from the workforce (conditional on surviving to 1950). We similarly re-weight these samples to resemble our population of interest (see columns 5 to 7 in Appendix Table A.1 for summary statistics).

3.2 1965 Dictionary of Occupational Titles

To explore the potential role of occupational task content in mediating the effect of Social Security on old-age mortality, we digitized data from the Third Edition of the U.S. Department of Labor's *Dictionary of Occupational Titles* (U.S. Department of Labor, 1965). More specifically, we digitized data from a 1966 supplement to the 1965 *Dictionary of Occupational Titles* (DOT), which contains detailed data on physical demands and working conditions for 13,815 unique job titles (U.S. Department of Labor, 1966).[7] Appendix Figure A.3 shows a few examples. The first column lists the DOT code associated with any given job title, which is listed in the fourth column (there are 3,133 unique DOT codes). For example, the DOT code "505.884" contains 5 job titles broadly

[6] For more details on Numident claims records and how linkages to the 1930 Census were made, see Fetter, Lockwood and Mohnen (2026).

[7] We used generative AI tools to convert the raw PDFs into excel tables and manually verified the output, correcting a handful of errors in the raw data during the process.

related to metal spraying, coating, and related work. The six columns under "Physical Demands" capture various physical requirements of the job, including strength requirements and whether the job requires various abilities such as climbing, crawling, talking, hearing, etc. The seven columns under "Working conditions" capture environmental characteristics of the job, such as whether it is typically performed indoors or outdoors (or both), and whether workers are exposed to various conditions such as extreme heat or cold, humidity, noise, hazards, fumes, etc. Appendix Figure A.4 provides definitions for each of the physical demands and working conditions, and Appendix Table A.3 shows the distribution of these task content measures by broad occupational category.

Since the 1950 census occupation codes we leverage in the 1930 Census are much coarser (269 unique codes), DOT job titles have to be mapped to census occupation codes, ideally with a set of weights that can be used to calculate average task content measures that place more weight on more common job titles. To that end, we leverage a digitized version of the 1965 *Dictionary of Occupational Titles* by Garro-Marín et al. (2021), which provides a mapping between DOT codes and 1970 census occupation codes with employment weights based on an augmented version of the April 1971 Current Population Survey (CPS) (National Academy of Sciences, 2006). To map 1970 census occupation codes to 1950 codes (with employment weights), we use data from the 1970 Census (Ruggles et al., 2025b). With these two crosswalks in hand, we construct the final crosswalk in three steps: (1) calculate the unweighted share of each task content measure across DOT job titles for each DOT code, (2) calculate the employment-weighted share of each task content measure across DOT codes for each 1970 census occupation code, and (3) calculate the employment-weighted share of each task content measure across 1970 census occupation codes for each 1950 census occupation code.[8]

4 Empirical Strategy

4.1 Comparing similar workers with differing Social Security eligibility

Our analysis follows the approach developed in Fetter, Lockwood and Mohnen (2026), which is based on the expansion in the definition of covered employment over time and its interaction with a worker's birth cohort. As we note above, only certain types of employment were covered under the 1935 Act and 1939 Amendments. Workers who worked in types of employment that were

[8]Since the augmented 1971 April CPS only contains 60,441 workers, 40 percent of DOT codes in the 1966 supplement are missing in the crosswalk from Garro-Marín et al. (2021). That said, these DOT codes represent only 17 percent of job titles in the 1966 supplement and are likely uncommon jobs that are hard to identify in household survey data. A handful of 1970 census occupation codes are also missing from the crosswalk (mainly residual occupation categories). Therefore, we are implicitly assigning a weight of zero to these missing codes, and our final crosswalk has missing task content information for a handful of 1950 census occupation codes (affecting less than 0.1 percent of individuals in our linked sample).

not covered until the 1950 Amendments—notably self-employment, domestic service, agriculture, nonprofit organizations, and state and local government—could not begin accumulating quarters of coverage until their employment was brought into the program. Using the eligibility rules from the 1935 Act and the 1939 and 1950 Amendments, we compute, for each combination of birth cohort and employment type (either 1935-covered or 1950-covered), the earliest age at which a worker who remains in the same type of employment could become permanently fully insured and thus eligible for retirement benefits at age 65.[9]

Figure 1 illustrates the resulting variation. Panel (a) plots the earliest age at which a worker from a given year and quarter of birth would become permanently insured, separately for those in employment covered from 1935 and those in employment first covered in 1950. Because 1950-covered workers could not start accruing quarters until 1951—14 years after 1935-covered workers began doing so—they reach eligibility at an older age. Panel (b) plots the earliest age at which workers could actually begin receiving monthly benefits, which illustrates the differential "bite" of later employment coverage by birth cohort. Workers in 1950-covered employment from later birth cohorts became eligible at younger ages, and since benefits could not be collected before age 65 regardless of coverage status, the effective gap in benefit receipt was smaller for later-born cohorts than for earlier-born cohorts.[10] We define our key measure of treatment, "Social Security exposure," as the additive inverse of this earliest eligibility age ($-1\times$ earliest eligibility age), so larger values correspond to earlier access to benefits.

We make use of this variation with information recorded in the 1930 Census, in which we observe an individual's employment five years before the Social Security Act was passed. For each worker, the Census recorded an occupation, industry, and class of worker (e.g., wage earner, salaried employee, or self-employed). We classify each combination of these three variables into the year when that type of employment first became covered by Social Security, drawing on the legislative history described in Section 2.[11] Some categories of wage and salary employment in the Social Security Act, such as work in private households or agricultural wage work, are relatively easily identified in Census industry and occupation codes. These codes do not directly identify work in nonprofit organizations or work for state and local governments other than in public administration, but hospitals, educational institutions, and welfare and religious organizations were the industries with the heaviest concentration of nonprofit and public-sector employment, and we

[9] "Permanently fully insured" means that even if a worker left covered employment, they would be eligible to receive benefits upon reaching age 65. In this discussion, we use the term "permanently insured" interchangeably.

[10] Early claiming—beginning as early as age 62—was introduced for men in 1961, but this is relevant only for cohorts born in 1897 or later, which are not part of our analysis.

[11] Fetter, Lockwood and Mohnen (2026) provide details on the classification scheme and a validation exercise showing that in the 1940 Census, our coding is highly predictive of workers reporting possession of a Social Security Number.

treat wage and salary workers in these sectors as covered beginning in 1950.[12]

As in Fetter, Lockwood and Mohnen (2026), the key source of identifying variation is that the industry-based exclusions in the original act created differences in coverage timing among workers performing very similar jobs. A janitor employed by a manufacturing firm, for instance, would have been covered from 1935, while a janitor working for a public school, a nonprofit hospital, or a religious organization would not have been covered until at least 1950. The differing treatment of workers performing similar tasks was noted at the time: Arthur J. Altmeyer, the Chairman of the Social Security Board, noted in 1944 that many employees of nonprofit institutions were not covered, despite that their "skills, tasks, and earnings... do not usually distinguish them from comparable employees in commerce or industry" (Altmeyer, 1944). Our preferred specification exploits this within-occupation, cross-industry variation by including occupation-by-cohort fixed effects, which absorb any average differences across occupations and cohorts and isolate the component of eligibility variation arising from employment in different industries within the same occupation.

To implement this idea, we exclude self-employed workers from our sample and estimate specifications of the form

$$y_i = \alpha_{b(i)o(i)} + \delta_{g(i)} + \beta \cdot (\text{Social Security exposure})_{b(i)g(i)} + \Lambda' \mathbf{x}_i + \varepsilon_i, \tag{1}$$

where y is an outcome for individual i, $\alpha_{b(i)o(i)}$ are fixed effects for the interaction of birth cohort b and three-digit occupation in 1930 o, $\delta_{g(i)}$ are fixed effects for Social Security coverage category g, and $\mathbf{x}_i$ is a vector of controls including fixed effects for race and state or country of birth. The three coverage categories in $\delta_{g(i)}$ are: covered in 1935, covered in 1950 via domestic work, and covered in 1950 via work in the nonprofit or state and local government sector. When analyzing the outcomes of wives, whose eligibility at the time depended, for the most part, on their husbands' work records, these variables are for their husbands, and we additionally include fixed effects for wives' own birth cohort. The coefficient of interest, β, captures the effect of a one-year increase in Social Security exposure—that is, of being predicted to become eligible one year earlier. Observations are weighted using inverse propensity score weights and standard errors are clustered at the three-digit industry level.

For interpreting our estimates, it is worth noting that some workers we classify as not covered until 1950—state and local government employees in particular—may have had access to pre-existing retirement systems, which is likely to attenuate our estimates. However, as noted in Fetter,

[12] In the Census, we cannot distinguish between employees of nonprofit organizations and those of state or local governments within these industries, nor can we observe whether government workers were already enrolled in an existing retirement system. We therefore treat all such workers as first covered in 1950. A small share of workers in these industries worked in for-profit firms, but any resulting misclassification would tend to attenuate our estimates.

Lockwood and Mohnen (2026), the types of government employees who contribute most to our comparisons, such as custodial and maintenance workers, were generally less likely to be enrolled in such plans than police, firefighters, or teachers, who either do not appear in the private sector or have little bearing on our estimates.[13] Among nonprofit organizations, some private colleges and universities had retirement plans by the 1940s, but nonprofessional staff were typically ineligible, and only about 8 percent of nonprofit hospitals offered pension plans at the start of the 1940s. Moreover, these pre-existing retirement systems typically differed from Social Security in important respects—most notably, they tended to provide less generous or no provision for survivor's benefits.

The occupations at the heart of our comparison are those present in both the for-profit sector and either the nonprofit or public sector, and to a lesser extent private households. Appendix Table A.2 displays the 20 largest occupation categories in the population in which there is meaningful variation in Social Security coverage status. These range from various categories of laborers, managers, and clerical workers that span many industries to more specific occupations such as carpenters, truck drivers, and janitors.

4.2 Validity checks

Our empirical strategy rests on the assumptions that the identifying variation is predictive of subsequent Social Security eligibility and that it is uncorrelated with other determinants of longevity. We present evidence validating these assumptions.

Predicted Social Security eligibility and subsequent claiming behavior. We first verify that our parameterization of Social Security eligibility is predictive of actual claiming behavior. The discussion above suggests that workers with greater Social Security "exposure" should be more likely to ever claim Social Security and, conditional on claiming, to do so at earlier ages. Table 1 reports estimates from equation (1) using our claims linked sample. In column (1), the sample is the 1930 population of men born in 1875–1888 and covered in 1935 or 1950 (excluding the self-employed), and the dependent variable is an indicator for being linked to a claims record. Note that some—in fact, the majority of—individuals in this population who claimed Social Security fail to be linked to the claims data, or simply do not appear in the publicly-available version, but the true likelihood of claiming Social Security should be positively correlated with the probability of being linked to a claims record. In column (2), we limit to the subset of men linked to a claims record and test whether our measure of Social Security exposure is related to the age at which they claimed Social Security benefits.

[13] Firefighters and police are excluded entirely; teachers are included, but very few are classified as being outside of the public or nonprofit sector.

The result in column (1) shows that eligibility one year earlier leads to a 0.3 percentage point increase in the likelihood of being linked to a claims record, relative to a baseline mean of 8 percent. Column 2 shows that, conditional on being linked to a claims record, eligibility one year earlier leads to claiming monthly benefits 0.11 years earlier, relative to a mean claiming age of 67.5. Both coefficient estimates are highly statistically significant. Figure 2 provides a more detailed view of these patterns, combining both the extensive and intensive margin predictions. In particular, we estimate equation (1) with dependent variables $\mathbf{1}(\text{linked to claim at age } a)$, for ages a ranging from 65 to 95, and $\mathbf{1}(\text{linked to claim made in year } t)$, for t ranging from 1940 to 1970. The results are strongly consistent with the predictions that underlie our empirical strategy. In particular, panel (a) shows that men with greater measured Social Security exposure have a greater likelihood of claiming, with the largest effects at younger ages (between 65 and 70). Moreover, since men working only in 1950-covered employment would have started to become eligible starting in 1952, it is consistent with our predictions that those who were eligible earlier had a higher likelihood of filing retirement claims during the 1940s, but a *lower* likelihood of filing claims in years from 1952 through 1954.

We do not formally treat the results in Table 1 or Figure 2 as a first stage—and in particular, we do not use them to scale our reduced-form estimates. As we noted above, only a subset of claims are linked to the 1930 Census, so these estimates tend to understate the true difference in probability of claiming. Nevertheless, they do establish that there is a strong relationship between our parameterization of Social Security eligibility and subsequent claiming behavior, even when restricting comparisons to fine occupational categories.

Selection into the Family Tree sample. Because only a subset of the population observed in the 1930 Census is connected to a profile on FamilySearch with a death year, we also test whether our Social Security exposure measure is correlated with the probability of appearing in our Family Tree sample. Selection into the sample could threaten internal validity if it is correlated with the variation we exploit. Encouragingly, Table 2 shows that men's Social Security exposure is not significantly related to the probability of appearing in our sample, for either men or their wives.

Placebo checks using 1930 characteristics. A final check on whether our approach allows for valid causal comparisons is to test whether our variation is correlated with household characteristics in 1930, five years prior to the introduction of Social Security. Table 3 reports the results. In both the full population (panel a) and our death-census linked sample (panel b), we find no evidence that Social Security exposure is systematically related to pre-existing characteristics such as home ownership, house values, radio ownership, living in an urban area, living on farm, or literacy.

5 Results

5.1 The impact of Social Security on old-age mortality

Table 4 presents our main regression results. The first column in panel (a) shows the impact of predicted Social Security eligibility on men's age at death. It implies that a 10-year earlier predicted eligibility—roughly the difference between being covered in 1935 vs. 1950 for the 1875 birth cohort—leads to a marginally-significant 0.31-year increase in age at death. Note that here we are implicitly conditioning on surviving to 1930, which entails surviving to different ages for different cohorts (from 42 for the 1888 birth cohort to 55 for the 1875 birth cohort). To make the outcomes more comparable across cohorts, in the second column we examine the impact of Social Security on age at death conditional on surviving to age 55 (which 93 percent of men in our sample do).[14] This shows that a 10-year earlier eligibility leads to a 0.45-year gain in age at death conditional on reaching age 55, which is significant at the 1 percent level. Column 4 tests whether earlier eligibility for Social Security is predictive of surviving to age 55 and reveals a precise null effect, alleviating concerns that the impact on conditional age at death is driven by positive selection. Therefore, Social Security had a meaningful impact on the longevity of early male recipients.

What about their wives? While few of them participated in the labor force, they could have indirectly benefited from their husbands' Social Security benefits as well as additional spousal benefits once they reached age 65 and potentially survivor's benefits, if they outlived their husbands. Results for wives of men in our sample are shown in panel (b) of Table 4. In addition to age at death conditional on both members of the couple surviving to 1930 (column 1), we also examine effects on age at death conditional on reaching age 55 (column 2) and age at death conditional on husbands reaching age 55 (column 3). Regardless of how we define outcomes, we find no impact of Social Security on wives' longevity, including the probability of reaching age 55 or husbands reaching age 55 (columns 4 and 5).

Figure 3 shows a visual representation of this asymmetry between men and wives. It plots average age at death conditional on surviving to age 55 for men and age at death conditional on husbands surviving to age 55 for wives by husbands' birth cohort, separately for husbands covered in 1935 vs. 1950 (where we have normalized conditional age at death relative to the 1888 birth cohort within each category). The patterns for men in panel (a) show that the difference in conditional age at death between workers covered earlier vs. later was greater for earlier cohorts, consistent with greater differences in predicted Social Security eligibility age as in panel (b) of Figure 1. Moreover, because the cohort gradient for later-covered workers is steeper than the

[14] Snyder and Evans (2006) and Noghanibehambari and Fletcher (2025) similarly examine mortality rates for a fixed age window when comparing different birth cohorts.

corresponding gradient for earlier-covered workers, this implies that the impact of earlier eligibility is positive. In contrast, the patterns for women in panel (b) reveal similar cohort gradients for the wives of 1935 vs. 1950-covered husbands, illustrating why we find no significant effect in the regression results.

To understand whether this gender asymmetry is due to differences in the nature of the treatment for husbands vs. wives, in panel (c) of Table 4 we examine impacts for women born in 1875–1888 who were never married and had non-missing employment information in 1930, and for whom we have year of death information from the Family Tree.[15] Since most women were married and did not participate in the labor force in this time period, this sample only includes around 64,000 women. Interestingly, we find no significant impact on these working women either, suggesting that other differences between men and women likely account for the lack of effects on women, a point we return to in the next section.

Table 5 presents separate estimates by race. For our preferred outcome of age at death conditional on surviving to age 55, we find significant effects for both White and Black men, but the magnitude of the coefficient is five times higher for Black men. A 10-year earlier predicted eligibility for Social Security led to a longevity gain of 1.64 years for Black men, relative to 0.32-year gain for White men. Similar to our baseline results, Social Security did not affect men's probability of reaching age 55 and we generally find no significant impact on wives' longevity for both race groups (with the exception of a marginally significant negative effect on Black women's probability of reaching age 55).

Appendix Figure A.5 shows that our main results, both the baseline estimates pooling everyone and the estimates by race, are robust to a variety of sensitivity checks, including using a dollar-based measure of exposure to Social Security and including additional control variables (3-digit industry fixed effects, the interaction of 3-digit occupation and Social Security coverage category fixed effects, controls for 1930 household characteristics, and 1930 county of residence fixed effects). That our estimates are robust to the inclusion of industry and occupation-by-coverage fixed effects highlights that our results are not driven by average differences between men covered in earlier vs. later, but rather by differential exposure to Social Security based on the interaction of cohort and coverage status. The county fixed effects alleviate the concern that our findings might reflect differences in characteristics of 1930 locations that are correlated with our Social Security eligibility variation. Among other things, this rules out differential place-based exposure to the Great Depression.

[15]The rules for Social Security eligibility for women were the same as for men until 1956, when early retirement at age 62 was introduced for women only (at a reduced benefits rate). This option was extended to men in 1961. Men and women in our cohorts of interest were not affected by these provisions.

5.2 Evidence on earlier retirement from physically-demanding or hazardous jobs

In this section, we explore one potential channel through which earlier eligibility for Social Security may have extended the lives of older men: earlier retirement from jobs that are physically demanding or involve poor working conditions. Manual work was much more common a century ago than it is today, even at older ages. For example, 68 percent of male workers aged 55 or older were farmers, blue-collar workers, or laborers in 1930, compared to just 34 percent in 2024. Working environments were also considerably riskier, as evidenced by significantly higher workplace fatalities (Centers for Disease Control and Prevention, 1999). Since past research suggests that manual work has adverse effects on health outcomes (Case and Deaton, 2005) and that the transition toward white-collar work over the 20th century has improved the health outcomes of older men in the U.S. (Costa, 2000, 2015), Social Security-induced earlier retirement from this type of work could plausibly have improved older men's longevity. Two aspects of our findings further support this hypothesis. First, that Social Security does not appear to have reduced the mortality of wives of early recipients suggests that the employment margin may have played an important role since female labor force participation was relatively low at the time. Second, that we find larger effects for Black men than White men, and no effects on never-married working women, suggests that the nature of jobs could be important, since Black men were more likely to hold manual jobs than White men and few women held such jobs.[16]

We start by documenting that earlier eligibility for Social Security induced earlier exits from the workforce among men. Table 6 shows the impact of Social Security eligibility on employment and labor force participation in 1950 using our sample of men linked across the 1930 and 1950 Censuses. The estimates in panel (a) pooling all men imply that being predicted to be eligible for Social Security 10 years earlier is associated with a 2-percentage-point lower probability of being employed or in the labor force in 1950. The estimates by race in panels (b) and (c) imply that the impact was twice as large for Black men than White men. Therefore, earlier eligibility for Social Security not only led to earlier Social Security claiming as we showed in Section 4.2, but also earlier exits from the workforce. These findings are consistent with research showing that the introduction of the Social Security early eligibility age of 62 in 1961 led to earlier claiming and lower earnings (Engelhardt et al., 2022).

Next, we show that both the longevity improvements and the labor-force exits were concentrated among men who held strenuous or hazardous jobs in 1930. Using the occupational task content data from the 1965 *Dictionary of Occupational Titles* described in Section 3.2, we divide jobs into two categories: (1) jobs that are physically demanding or in which workers are exposed

[16] In our sample, the share of Black men engaged in manual work in 1930 (farmers, operatives, craftsmen, laborers) is 76 percent, compared to 70 percent for White men and just 17 percent for never-married working women. As we will show, this modest racial gap masks larger differences in the type of manual work performed.

to poor working conditions, and (2) all other jobs. We define occupations in (1) as those that are in the top quartile based on the share of corresponding DOT job titles that are "very" physically demanding, in the sense that they can require lifting or carrying objects of over 100 lbs. (and frequently lifting or carrying objects of 50 lbs. or more), or in which workers are exposed to "hazards" or "toxic" working conditions (fumes, odors, toxic conditions, dust, poor ventilation). Appendix Table A.4 lists the subset of those occupations with at least 1,000 workers in our population of interest and with some variation in Social Security coverage. Occupations that are very physically demanding include foresters, stationary firemen, and lumbermen. Occupations in which workers are exposed to hazards include roofers, sheet metal workers, and boilermakers. Occupations in which workers are exposed to toxic working conditions include welders, painters, and furnacemen. More common occupations that fall in the top quartile include carpenters and laborers.

Panel (a) in Table 7 shows estimates of the impact of Social Security on age at death conditional on surviving to age 55, separately by occupation group. Note that while strenuous or hazardous jobs make up just a quarter of occupations, they employ around 45 percent of workers in our sample. We find that earlier Social Security eligibility had a significant positive effect on the conditional age at death of workers in jobs that are physically demanding or involve poor working conditions, and no significant effect on other workers. Moreover, the estimates in panel (b) show that earlier exits from the workforce are also concentrated among the former group, consistent with the notion that increased longevity is tied to relief from hard labor and reduced exposure to harsh working conditions.

Table 7 also presents separate estimates for White and Black men, revealing similar patterns for both groups. Black men were much more likely than White men to hold physically-demanding or hazardous jobs (66 percent vs. 45 percent), which likely accounts for part of the racial gap in the baseline effect on old-age mortality. That said, the estimates in panel (a) show an even larger racial gap among workers in physically-demanding or hazardous jobs, with a 10-year earlier eligibility for Social Security raising Black men's conditional age at death by 2.3 years vs. 0.4 years for White men. What could account for this? One possible explanation is that Black workers were concentrated in worse jobs than their White counterparts. For example, Foote et al. (2003) find that in the early 20th century Black workers at the Ford Motor Company were disproportionately assigned to hot, dangerous foundry jobs. Moreover, contemporary accounts suggest that Black workers were often assigned to less desirable tasks than their White co-workers, even conditional on having the same job title (Spero and Harris, 1931).

Overall, these findings suggest that earlier Social Security eligibility led men to retire earlier from jobs in which continued employment tended to have negative health effects, ultimately extending their longevity. This channel fits the patterns for men vs. women and White vs. Black men,

given differences in participation in physically demanding or hazardous jobs at the time.[17] While past studies on the impact of Social Security on old-age mortality have documented evidence suggesting instead that *prolonged* employment is beneficial for longevity (Snyder and Evans, 2006; Fitzpatrick and Moore, 2018), we highlight a few key differences. First, these studies rely on different sources of variation. Snyder and Evans (2006) leverage a discontinuity in Social Security benefit amounts between the 1916 and 1917 birth cohorts due to the Social Security "Notch," while Fitzpatrick and Moore (2018) compare people just below and above the early Social Security eligibility age of 62. Second, both of these studies focus on short-run mortality rates, whereas we focus on completed longevity. Lastly, the time periods we study are very different. Whereas our cohorts worked at a time when manual work was pervasive and working conditions were poor relative to today, Snyder and Evans (2006) and Fitzpatrick and Moore (2018) focus on later cohorts who were more likely to be white-collar workers and enjoy better working conditions. Our findings are consistent with two studies in the context of Israel and Spain who find that delayed retirements due to pension reforms had negative effects on longevity for individuals in blue-collar or low-skill jobs (Saporta-Eksten et al., 2021; Bellés-Obrero et al., 2025).

6 Conclusion

The effect of Social Security on longevity is important for understanding the welfare, distributional, and fiscal consequences of the program. In this paper, we provide new evidence on the causal impact of the early Social Security program on the mortality of recipients and their spouses. Combining an empirical approach that exploits within-occupation, cross-industry differences in Social Security coverage with a dataset connecting death information from from profiles on FamilySearch to the 1930 Census, we find that men who were more likely to have received Social Security, or who received it earlier, lived longer as a result. In contrast, we find no evidence that the early Social Security program affected the longevity of recipients' wives or of unmarried working women.

Preliminary evidence on the mechanisms underlying these findings points to earlier exit from physically demanding or hazardous jobs as an important channel. Both the longevity effects and the labor force participation effects of Social Security are concentrated among men in occupations involving heavy physical demands or exposure to poor working conditions, while the effects for men in other occupations are small and insignificant. This channel is also consistent with the asymmetry between the effects on men and wives, since the health consequences of physically-demanding work are borne directly by the worker.

[17] Only 2.3 percent of never-married working women in our sample held such jobs, which could explain why we fail to find any longevity gains for this group.

Our findings have several implications. First, they suggest that the fiscal cost of Social Security may be somewhat higher than calculations that treat mortality as exogenous would imply, since the program extended recipients' lives—and hence the duration of benefit receipt. Second, the concentration of longevity gains among workers in physically demanding and hazardous occupations suggests that the distributional effects of Social Security may have been more progressive than they otherwise would have been, since these workers tend to be at the lower end of the earnings distribution. Third, that the longevity effects of Social Security appear to depend critically on the nature of recipients' work suggests that as the composition of the labor force has shifted away from physically demanding occupations over the twentieth century, the longevity effects of Social Security for the marginal retiree may have changed as well. This may help reconcile our findings with more recent evidence suggesting that retirement can be harmful to health in some settings.

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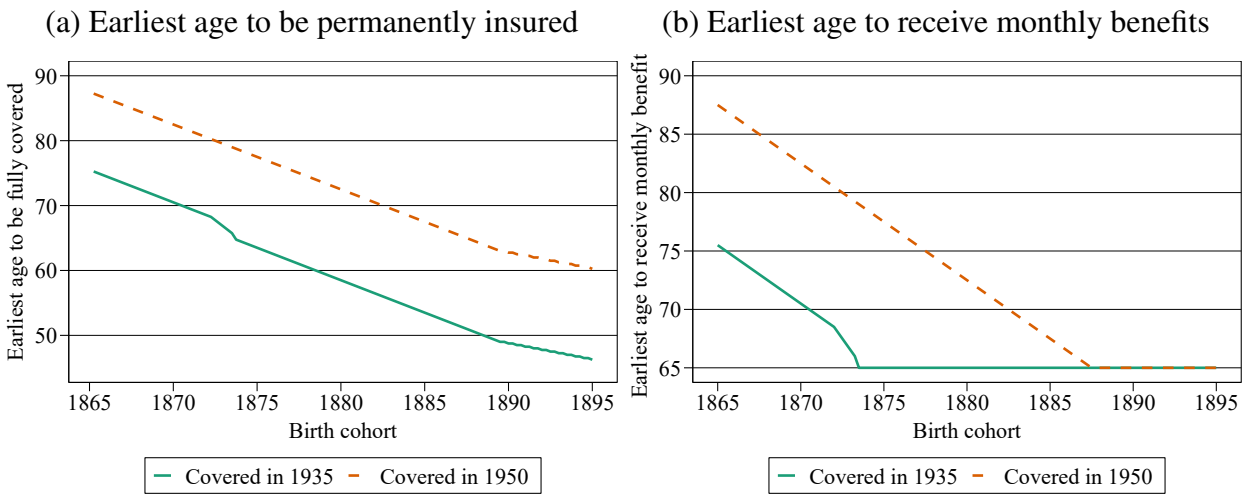
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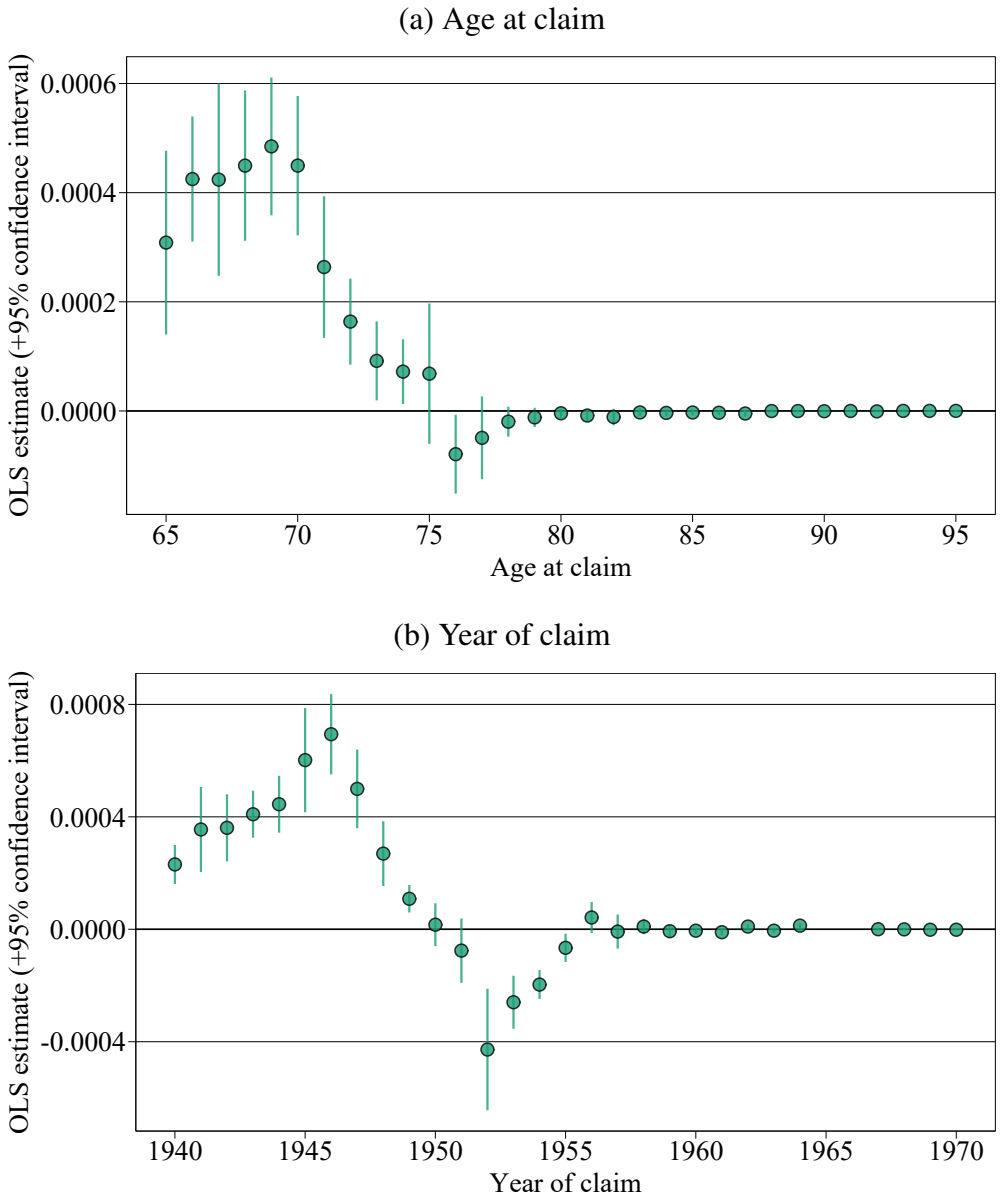
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Figure 1: Earliest age to be permanently insured and earliest age to receive monthly Social Security benefits



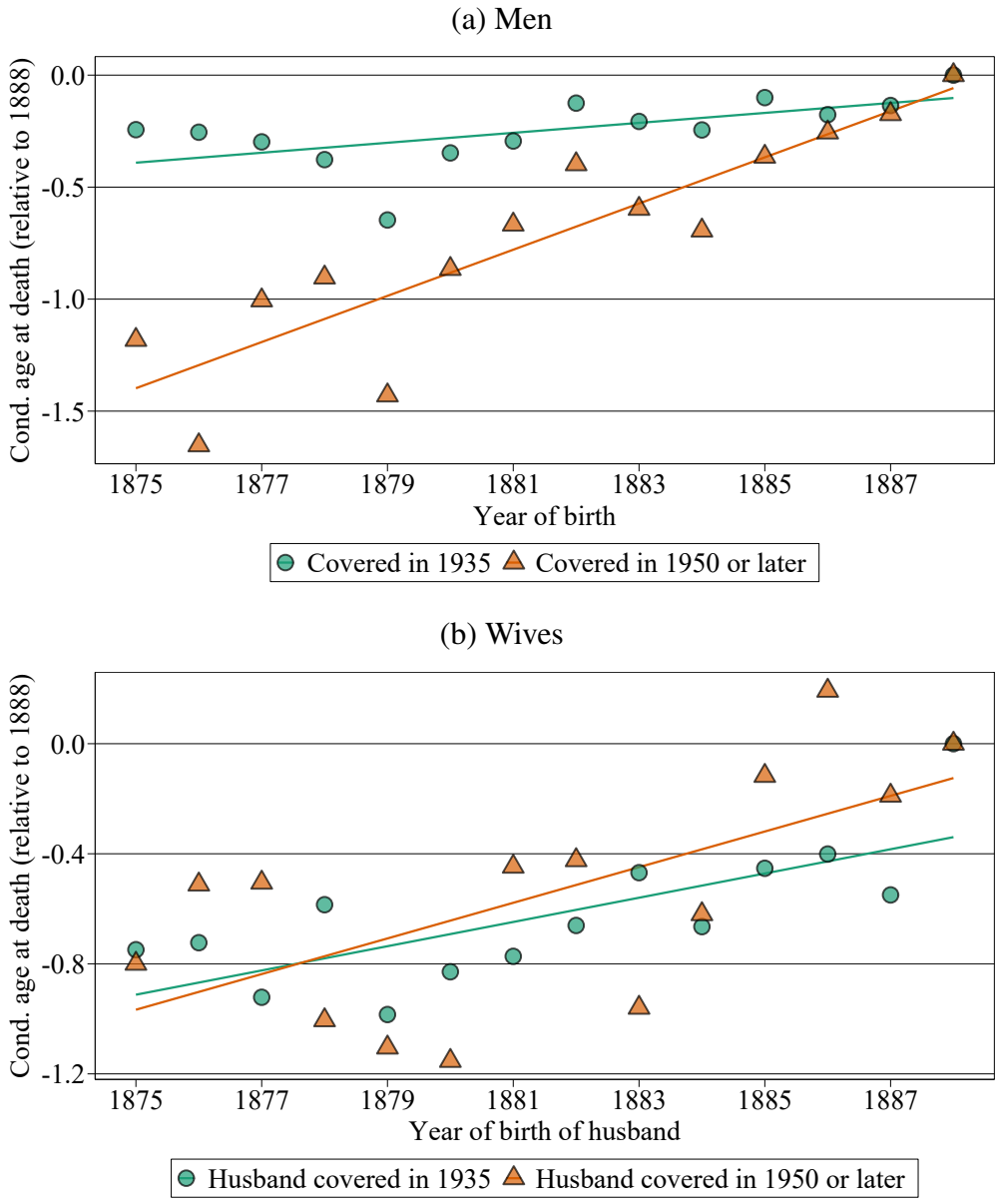
Notes: Minimum ages are based on Social Security Act of 1935, 1939 Amendments, and 1950 Amendments, assuming continuous employment in 1935 or 1950-covered employment. The minimum ages shown are those under the 1950 Amendments.

Figure 2: The impact of exposure to Social Security on age at claim and year of claim



Notes: Panel (a) shows the coefficients on Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1), where the sample consists of all men in our population of interest in the 1930 Census and the dependent variable is an indicator for appearing in our claims linked sample and claiming at a specified age. Panel (b) plots analogous estimates for claiming in a specified year. All regressions include fixed effects for men's 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace. 95% confidence intervals based on robust standard errors clustered by men's 3-digit industry. Claims made in 1965 and 1966 are excluded from the sample to avoid capturing Prouty beneficiaries.

Figure 3: Average conditional age at death by cohort and Social Security coverage status



Notes: Panel (a) plots the average age at death by birth cohort and Social Security coverage status among men in our linked sample who survived to age 55. Panel (b) plots the average age at death by husband's birth cohort and Social Security coverage status among wives in our linked sample whose husbands survived to age 55. Ages at death are normalized to be relative to the 1888 birth cohort (within each group). Observations weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930.

Table 1: The impact of Social Security eligibility on subsequent claiming behavior of men

	Dependent variable:	
	Linked to claims data	Age at claim
	(1)	(2)
Social Security exposure	0.0030***	-0.1087***
	(0.0004)	(0.0152)
Mean of dep. var.	0.08	67.52
N	5,028,910	421,959

Notes: This table shows the coefficients on Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1), where the sample consists of all men in our population of interest in the 1930 Census in column 1 and the subset of those men appearing in our claims linked sample in column 2. In column 1, the outcome is an indicator for appearing in our claims linked sample. In column 2, the outcome is age at the time of claiming. All regressions include fixed effects for men's 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace. In column 2, observations are weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930. Robust standard errors in parentheses, clustered by men's 3-digit industry. *** 1%, ** 5%, * 10% significance.

Table 2: Selection of 1930 individuals into Family Tree sample

	Dep. var: In Family Tree sample	
	Men	Wives
	(1)	(2)
Social Security exposure	0.0002	0.0001
	(0.0004)	(0.0003)
Mean of dep. var.	0.35	0.31
N	5,028,910	3,799,753

Notes: This table shows the coefficients on own or husband's Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1), where the sample consists of all men in our population of interest in the 1930 Census and the outcome is an indicator for appearing in our Family Tree sample. All regressions include fixed effects for men's 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace. Robust standard errors in parentheses, clustered by men's 3-digit industry. *** 1%, ** 5%, * 10% significance.

Table 3: Placebo checks: 1930 characteristics

	Dependent variable: Characteristic in 1930					
	Homeowner	House value (log)	Own a radio	Live in urban area	Live on farm	Literate
	(1)	(2)	(3)	(4)	(5)	(6)
Panel (a): Population						
Social Security exposure	-0.0005	0.0121	0.0003	0.0001	0.0001	-0.0004
	(0.0006)	(0.0082)	(0.0009)	(0.0007)	(0.0002)	(0.0002)
Mean of dep. var.	0.49	8.91	0.45	0.73	0.07	0.94
N	4,857,510	2,317,954	5,028,910	5,028,910	5,028,910	5,028,910
Panel (b): Men in Family Tree sample						
Social Security exposure	0.0015	-0.0003	0.0007	0.0003	-0.0000	-0.0010*
	(0.0075)	(0.0014)	(0.0008)	(0.0004)	(0.0004)	(0.0006)
Mean of dep. var.	0.49	8.67	0.46	0.73	0.07	0.94
N	1,730,285	897,020	1,743,367	1,743,367	1,743,367	1,743,367

Notes: This table shows the coefficients on own or husband's Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1), where the outcome is a specified characteristic (indicated in the column titles). In panel (a), the sample consists of all men in our population of interest in the 1930 Census while in panel (b) the sample only includes the subset of men who appear in our Family Tree sample. All regressions include fixed effects for men's 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace. In panel (b), observations are weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930. Robust standard errors in parentheses, clustered by men's 3-digit industry. *** 1%, ** 5%, * 10% significance.

Table 4: The impact of Social Security eligibility on old-age mortality

	Dependent variable:				
	Age at death			Survived to age 55	
	Full sample	Conditional on surviving to age 55	Conditional on husband surviving to age 55	Self	Husband
	(1)	(2)	(3)	(4)	(5)
Panel (a): Men					
Social Security exposure	0.031*	0.045***	—	0.000	—
	(0.017)	(0.016)		(0.001)	
Mean of dep. var.	72.8	74.5		0.93	
N	1,743,367	1,627,855		1,743,367	
Panel (b): Wives					
Social Security exposure	-0.002	-0.007	-0.016	0.000	-0.001
	(0.019)	(0.016)	(0.021)	(0.000)	(0.001)
Mean of dep. var.	76.3	78.7	76.4	0.92	0.94
N	1,193,431	1,107,147	1,121,049	1,193,431	1,193,431
Panel (c): Never-married working women					
Social Security exposure	0.010	-0.051	—	0.002	—
	(0.116)	(0.105)		(0.002)	
Mean of dep. var.	78.5	80		0.95	
N	64,245	61,209		64,245	

Notes: This table shows the coefficients on own or husband's Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1), where the sample consists of individuals in our Family Tree sample and the outcome is either unconditional age at death, age at death conditional on surviving to age 55 (or conditional on husbands surviving to age 55), or an indicator for surviving to age 55 (or husbands surviving to age 55). All regressions include fixed effects for 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace (or husbands' characteristics). The regressions for wives in panel (b) additionally include fixed effects for their own birth cohort. Observations weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930. Robust standard errors in parentheses, clustered by 3-digit industry (or husbands' 3-digit industry). *** 1%, ** 5%, * 10% significance.

Table 5: The impact of Social Security eligibility on old-age mortality: Heterogeneity by race

	Dependent variable:				
	Age at death			Survived until age 55	
	Full sample	Conditional on surviving to age 55	Conditional on husband surviving to age 55	Self	Husband
	(1)	(2)	(3)	(4)	(5)
Panel (a): White men					
Social Security exposure	0.019	0.032***	—	0.000	—
	(0.014)	(0.008)		(0.001)	
Mean of dep. var.	73	74.5		0.93	
N	1,716,622	1,603,697		1,716,622	
Panel (b): Black men					
Social Security exposure	0.155**	0.164**	—	0.001	—
	(0.063)	(0.078)		(0.002)	
Mean of dep. var.	71.1	73.6		0.89	
N	16,622	14,807		16,622	
Panel (c): Wives of white men					
Social Security exposure	0.011	0.010	0.008	0.000	0.000
	(0.016)	(0.010)	(0.019)	(0.000)	(0.001)
Mean of dep. var.	76.5	78.7	76.6	0.93	0.94
N	1,179,428	1,094,850	1,108,199	1,179,428	1,179,428
Panel (d): Wives of Black men					
Social Security exposure	-0.136	-0.104	-0.242	-0.004*	-0.002
	(0.134)	(0.161)	(0.150)	(0.003)	(0.004)
Mean of dep. var.	72.6	76.8	73	0.86	0.9
N	7,710	6,638	6,976	7,710	7,710

Notes: This table shows the coefficients on own or husband's Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1), where the sample consists of individuals in our Family Tree sample and the outcome is either unconditional age at death, age at death conditional on surviving to age 55 (or conditional on husbands surviving to age 55), or an indicator for surviving to age 55 (or husbands surviving to age 55). Panels (a) and (b) present estimates for White and Black men, while panels (c) and (d) present estimates for their wives. All regressions include fixed effects for 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace (or husbands' characteristics). The regressions for wives in panels (c) and (d) additionally include fixed effects for their own birth cohort. Observations weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930. Robust standard errors in parentheses, clustered by 3-digit industry (or husbands' 3-digit industry). *** 1%, ** 5%, * 10% significance.

Table 6: The impact of Social Security eligibility on labor force attachment in 1950

	Dependent variable:	
	In labor force in 1950	Employed in 1950
	(1)	(2)
Panel (a): All men		
Social Security exposure	-0.0020**	-0.0020*
	(0.0009)	(0.0012)
Mean of dep. var.	0.61	0.57
N	1,826,487	1,826,487
Panel (b): White men		
Social Security exposure	-0.0020**	-0.0019*
	(0.0008)	(0.0011)
Mean of dep. var.	0.61	0.58
N	1,745,756	1,745,756
Panel (c): Black men		
Social Security exposure	-0.0041*	-0.0043**
	(0.0022)	(0.0021)
Mean of dep. var.	0.58	0.54
N	63,322	63,322

Notes: This table shows the coefficients on Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1), where the sample consists of men in the 1930–1950 Census linked sample and the outcome is either an indicator for being in the labor force in 1950 or an indicator for being employed in 1950. Panel (a) show results pooling all men while panels (b) and (c) separately look at white men and Black men. All regressions include fixed effects for 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace. Observations weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930. Robust standard errors in parentheses, clustered by 3-digit industry. *** 1%, ** 5%, * 10% significance.

Table 7: The impact of Social Security eligibility on old-age mortality and labor force participation in 1950: Workers in physically-demanding jobs or jobs with poor working conditions vs. all other workers

	Workers in physically-demanding jobs or jobs with poor working conditions			All other workers		
	All men	White men	Black men	All men	White men	Black men
	(1)	(2)	(3)	(4)	(5)	(6)
Panel (a): Dependent variable: Age at death (conditional on surviving to age 55)						
Social Security exposure	0.062**	0.040**	0.234**	0.020	0.017	0.033
	(0.025)	(0.017)	(0.118)	(0.025)	(0.018)	(0.109)
Mean of dep. var.	74.5	74.6	73.7	74.5	74.5	73.5
N	717,043	701,378	9,962	901,270	892,858	4,785
Panel (b): Dependent variable: In labor force in 1950						
Social Security exposure	-0.004***	-0.003***	-0.006***	-0.001	-0.001	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.005)
Mean of dep. var.	0.56	0.56	0.55	0.65	0.65	0.63
N	806,762	754,040	41,927	1,008,862	981,154	21,123

Notes: This table shows the coefficients on Social Security exposure (i.e., negative of earliest age of Social Security eligibility) from equation (1). In panel (a), the sample consists of men in our Family Tree sample and the outcome is age at death conditional on surviving to age 55. In panel (b), the sample consists of men in the 1930–1950 Census linked sample and the outcome in an indicator for being in the labor force in 1950. Columns 1–3 restrict the sample to men in physically-demanding or jobs with poor working conditions, while columns 4–6 restrict the sample to men in all other jobs (either pooling all men or restricting to white or Black men). All regressions include fixed effects for 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace. Observations weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930. Robust standard errors in parentheses, clustered by 3-digit industry. *** 1%, ** 5%, * 10% significance.

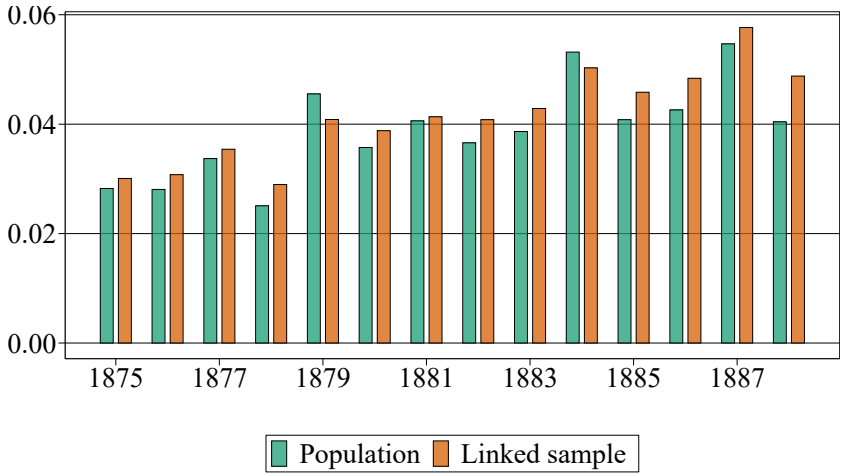
Online Appendix for

The Introduction of Social Security and Elderly Mortality

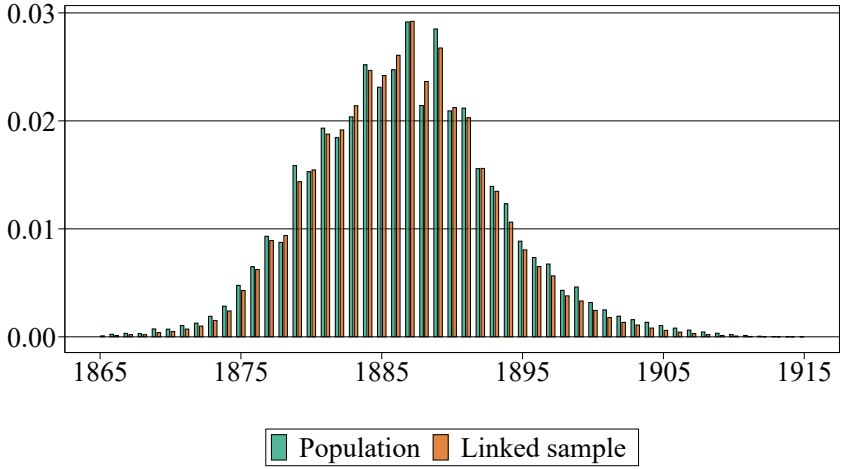
By Daniel K. Fetter, Lee M. Lockwood, Joseph Price, and Paul Mohnen

Figure A.1: Cohort coverage in Family Tree sample

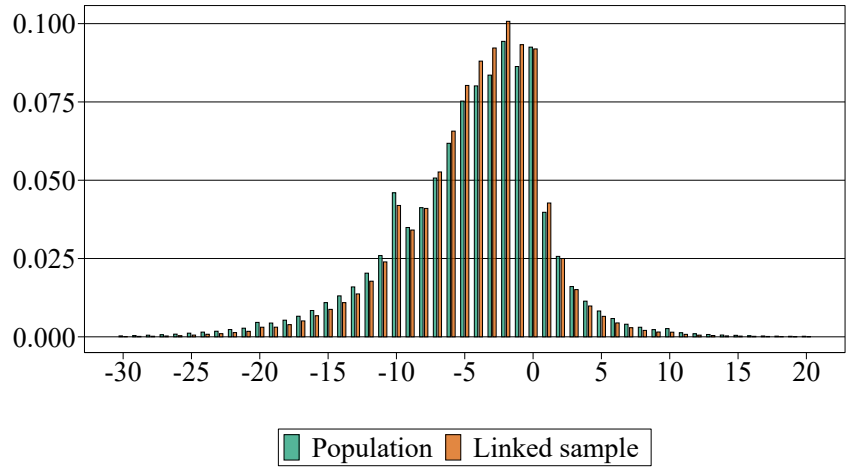
(a) Year of birth for men born in 1875–1888 and covered by Social Security in 1935 or 1950 (excluding self-the employed)



(b) Year of birth for wives of men in panel (a)

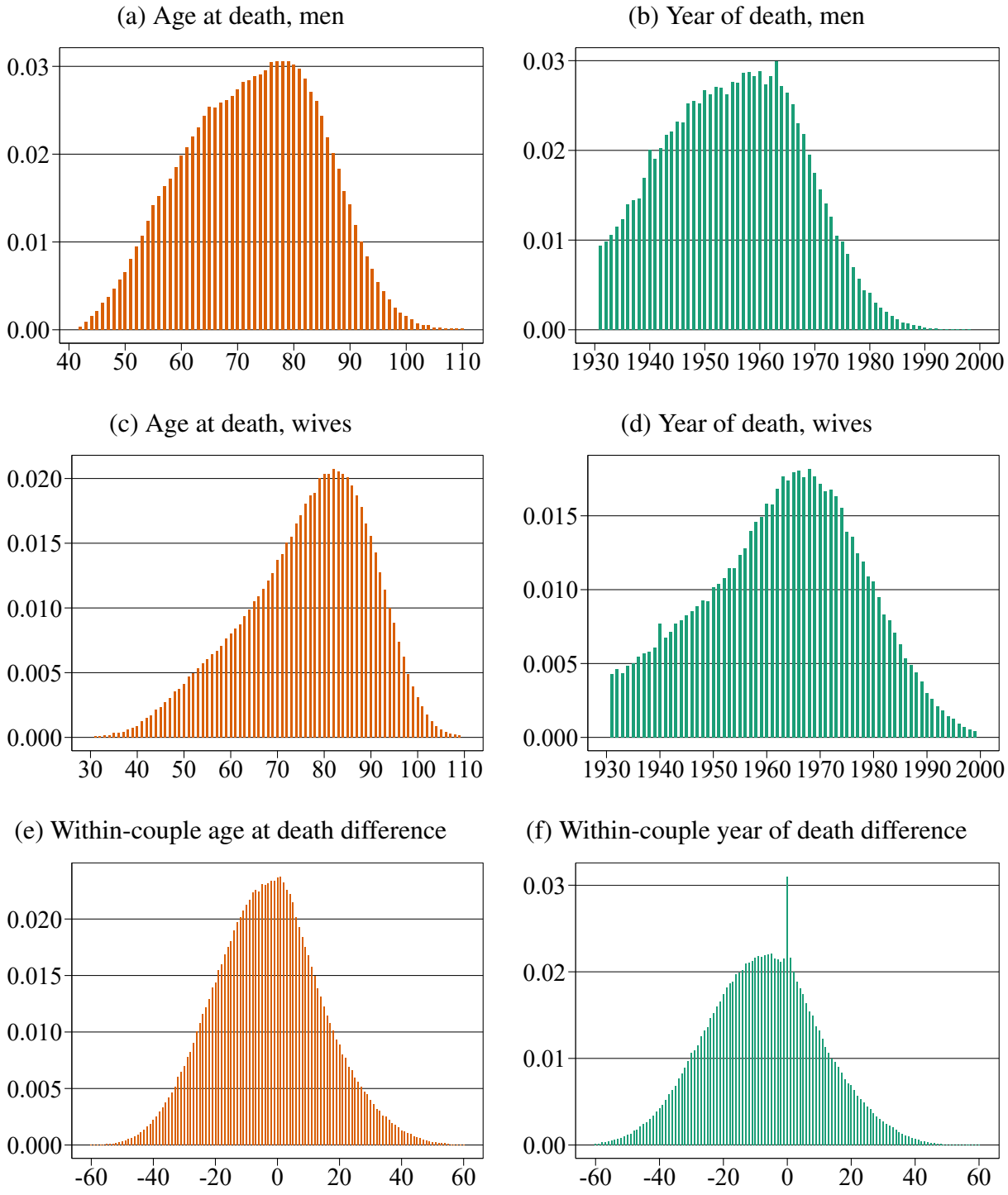


(c) Within-couple year of birth differences (husband minus wife)



Notes: Sample in panel (a): men born in 1875–1888 in the 1930 Census, covered by Social Security in 1935 or 1950 based on their employment information in 1930 (excluding the self-employed). Sample in panel (b): wives of men satisfying restrictions in panel (a). Sample in panel (c): couples in panels (a) and (b).

Figure A.2: Distribution of year of death and age at death in Family Tree sample



Notes: Panels (a) and (c) plot the age at death distribution among men and wives in our Family Tree sample. Panel (e) plots within-couple age at death differences (husband minus wife). Panels (b), (d) and (f) show analogous distributions for year of death. Observations are weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930.

Figure A.3: Examples from 1966 supplement to 1965 Dictionary of Occupational Titles

D.O.T. Code	WT Group Page No.	Industry	D.O.T. Title	Physical Demands						Working Conditions							Training Time	
				1	2	3	4	5	6	1	2	3	4	5	6	7	GED	SVP
505.782	435	Any Ind	Metal-Spraying-Machine Operator, Automatic I.	M			4		6	I		3		5	6	7	4	7
505.782	435	Any Ind	Metal-Spraying-Machine Operator, Crucible Gun.	M			4		6	I		3			6	7	4	7
505.782	435	Elec. Equip	Pasting-Machine Operator	L			4		6	I				5			3	4
505.782	435	Nonfer. Metal Alloys.	Alodize-Machine Operator	M			4		6	I			4				3	5
505.782	435	Tinware	Coater-Machine Operator	M	2		4		6	I						7	3	6
505.884	322	Any Ind	Electroless Plater	M			4			I							2	3
505.884	322	Any Ind	Metal Sprayer, Production	M			4		6	I				5			2	2
505.884	322	Any Ind	Solder Sprayer	M		3	4		6	I					6	7	3	3
505.884	322	Ordnance	Case Preparer-and-Liner Operator.	L			4		6	I					6		3	4
505.884	322	Plumb. Supplies	Enameler	L			4		6	I		3					3	6
505.885	447	Ammunition	Lubricating-Machine Tender	M			4			I		3					1	2
505.885	447	Any Ind	Ceramic Coater, Machine	M			4		6	I		3					2	2
505.885	447	Any Ind	Metal-Spraying-Machine Operator, Automatic II.	M			4			I				5			2	2
505.885	447	Any Ind	Vacuum-Metalizer Operator	M			4		6	I							2	3
505.885	447	Button	Lacquer-Dipping-Machine Operator.	M			4			I							2	2
505.885	447	Firearms	Browning Processor	M			4		6	I		3	4		6		2	3
505.885	447	Nonfer. Metal Alloys.	Alodize-Machine Helper	M		3	4			I			4				2	3
505.885	447	Tinware	Coater Operator	L			4			I				5			2	3
505.885	447	Tinware	Compound Mixer	V		3	4			B					6		3	4
505.886	356	Drug. Prep. & Rel. Prod.; Soap.	Wax-Spraying-Machine Operator.	L			4			I							1	2

Figure A.4: Legend for physical demands and working conditions from 1966 supplement to 1965 Dictionary of Occupational Titles

PHYSICAL DEMANDS

1—Strength (Lift, Carry, Push, Pull):

		Maximum Lift	*Frequent Lift/Carry*
S	Sedentary	10 lbs.	
L	Light	20 lbs.	Up to 10 lbs. [1]
M	Medium	50 lbs.	Up to 25 lbs.
H	Heavy	100 lbs.	Up to 50 lbs.
V	Very Heavy	Over 100 lbs.	Over 50 lbs.

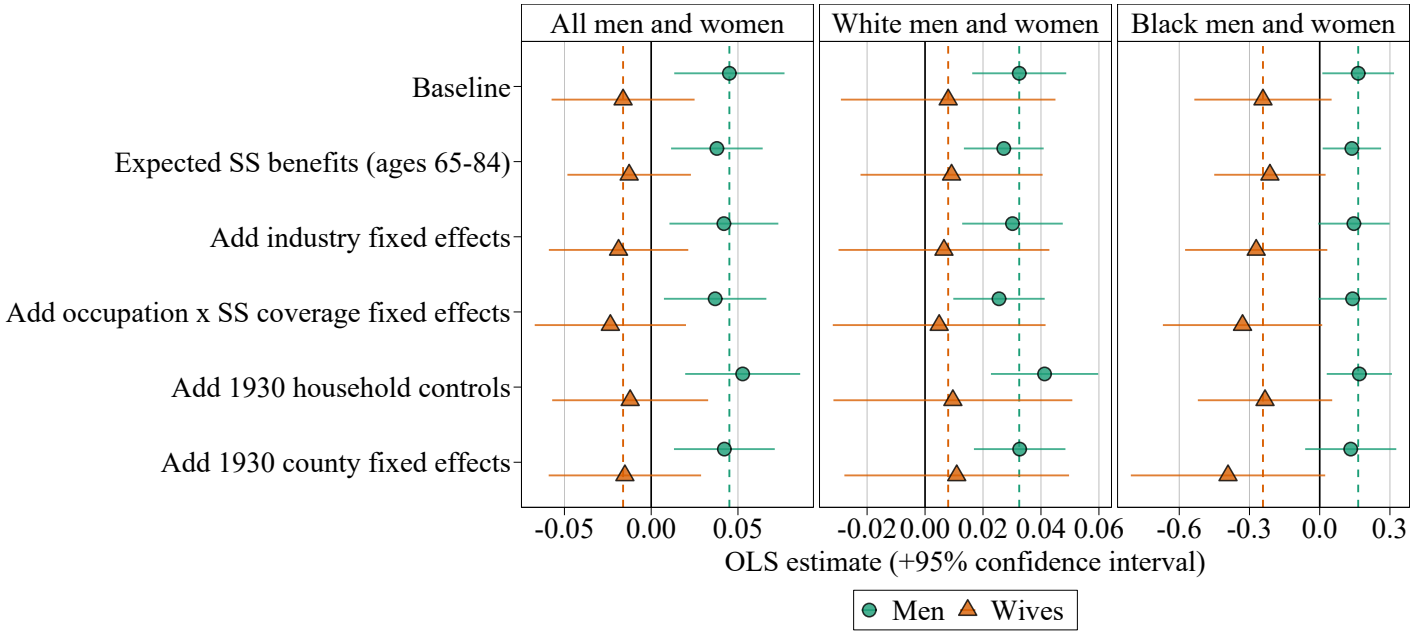
2—Climb and Balance
3—Stoop, Kneel, Crouch, Crawl
4—Reach, Handle, Finger, Feel
5—Talk, Hear
6—See (Acuity, Depth Perception, Field of Vision, Accommodation)

WORKING CONDITIONS

I Inside (75% or more)
O Outside (75% or more)
B Both

2—Extremes of Cold Plus Temperature Changes
3—Extremes of Heat Plus Temperature Changes
4—Wet and Humid
5—Noise and Vibration
6—Hazards
7—Fumes, Odors, Toxic Conditions, Dust, Poor Ventilation

Figure A.5: The impact of Social Security eligibility on old-age mortality: Robustness checks



Notes: The top row in this figure shows the baseline coefficients from Tables 4 and 5, where the outcome is age at death conditional on surviving to age 55 for men and age at death conditional on husbands surviving to age 55 for wives. The second row shows estimates using an alternative dollar-based measure of Social Security exposure (expected average monthly benefits between ages 65 and 84 based on the earnings history of a median earner, in tens of 2000 dollars). The bottom four rows show estimates under variants of the baseline specification that include various additional controls (3-digit industry fixed effects, 3-digit occupation-by-Social Security coverage category fixed effects, fixed effects for all outcomes in Table 3 except house values which is included as a linear control, or fixed effects for 1930 county of residence). Colors/shapes indicate separate estimates for men and wives. All regressions include fixed effects for men's or husbands' 3-digit occupation-by-birth cohort, Social Security coverage category, race, and birthplace. The regressions for wives additionally include fixed effects for their own birth cohort. Observations weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930. 95% confidence intervals based on robust standard errors clustered by men's or husbands' 3-digit industry.

Table A.1: Summary statistics in population vs. linked samples

		Family Tree sample		Claims linked sample		1930–1950 linked sample	
	Population	Unweighted	Weighted	Unweighted	Weighted	Unweighted	Weighted
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Black	0.078	0.01	0.079	0.027	0.081	0.035	0.077
Born in 1875–1879	0.295	0.286	0.296	0.351	0.285	0.238	0.294
Born in 1880–1884	0.376	0.369	0.376	0.394	0.375	0.376	0.376
Born in 1885–1888	0.328	0.345	0.328	0.216	0.34	0.387	0.33
Born abroad	0.307	0.142	0.306	0.262	0.309	0.286	0.306
Born in Northeast	0.207	0.217	0.209	0.245	0.206	0.219	0.207
Born in Midwest	0.25	0.378	0.254	0.312	0.248	0.294	0.251
Born in South	0.211	0.23	0.206	0.154	0.213	0.175	0.211
Born in West	0.025	0.033	0.025	0.027	0.025	0.026	0.025
Literate	0.94	0.976	0.942	0.97	0.939	0.959	0.941
Covered by SS in 1935	0.871	0.881	0.877	0.93	0.909	0.888	0.877
Covered by SS in 1950	0.129	0.119	0.123	0.07	0.091	0.112	0.123
Farmer	0.002	0.003	0.003	0.002	0.002	0.003	0.003
Unskilled worker	0.305	0.221	0.268	0.209	0.264	0.243	0.28
Skilled/semi-skilled worker	0.45	0.465	0.465	0.517	0.497	0.482	0.468
White-collar worker	0.243	0.311	0.264	0.273	0.237	0.273	0.25
Live in Northeast	0.364	0.257	0.316	0.373	0.361	0.363	0.365
Live in Midwest	0.322	0.385	0.352	0.369	0.335	0.355	0.328
Live in South	0.195	0.224	0.202	0.139	0.185	0.166	0.192
Live in West	0.118	0.134	0.13	0.119	0.118	0.117	0.115
Live on farm	0.066	0.077	0.07	0.043	0.068	0.063	0.067
Live in urban area	0.731	0.662	0.726	0.766	0.729	0.724	0.73
Own home	0.472	0.549	0.488	0.547	0.471	0.547	0.475
Own a radio	0.455	0.512	0.463	0.551	0.452	0.517	0.457
N	5,028,910	1,743,367		421,959		1,826,487	

Notes: This table shows mean characteristics of our population of interest in the 1930 Census (men born in 1875–1888 covered by Social Security in 1935 or 1950, excluding the self-employed) against corresponding unweighted and weighted means in different linked samples. Weighted means based on inverse propensity score weights that re-weight the linked samples to match the population in terms of cohort, race, birthplace, literacy status, farm status, urban status, homeonwership status, and radio ownership status.

Table A.2: Top 20 occupations with some variation in Social Security coverage year

	Summary statistics among men born in 1875–1888 in 1930 Census						
	Population			Family Tree linked sample		Claims linked sample	
Occupation	Number (1)	Share covered in 1935 (2)	Share (3)	Share covered in 1935 (4)	Share (5)	Share covered in 1935 (6)	Share (7)
Laborers (n.e.c.)	795,300	89.3	15.8	90.5	14.3	92.8	14.7
Operative and kindred workers (n.e.c.)	430,820	99.0	8.6	99.1	8.4	99.4	10.0
Salesmen and sales clerks (n.e.c.)	367,978	99.1	7.3	99.1	7.5	99.2	7.3
Managers, officials, and proprietors (n.e.c.)	325,231	98.2	6.5	98.0	7.4	98.7	6.4
Carpenters	245,313	98.5	4.9	98.5	5.5	98.7	5.4
Machinists	136,907	99.3	2.7	99.4	2.9	99.5	3.3
Foremen (n.e.c.)	132,982	98.3	2.6	98.5	3.1	99.1	3.1
Clerical and kindred workers (n.e.c.)	97,906	92.2	1.9	92.2	1.9	95.1	2.0
Painters, construction and maintenance	92,578	97.2	1.8	97.2	1.6	97.4	1.7
Truck and tractor drivers	82,226	95.5	1.6	95.1	1.8	96.6	1.8
Stationary engineers	73,678	92.5	1.5	92.8	1.6	95.2	1.6
Janitors and sextons	72,020	58.6	1.4	55.0	1.6	67.9	1.4
Insurance agents and brokers	51,911	99.8	1.0	99.7	1.2	99.9	1.0
Mechanics and repairmen, automobile	49,581	99.2	1.0	99.2	1.0	99.4	1.1
Plumbers and pipe fitters	48,628	96.8	1.0	96.9	1.0	97.6	1.0
Brickmasons, stonemasons, and tile setters	43,349	99.0	0.9	99.1	0.9	99.3	1.0
Cooks, except private household	42,331	95.8	0.8	95.0	0.5	95.6	0.7
Electricians	42,196	95.4	0.8	95.3	0.9	97.2	0.9
Tailors and tailoresses	36,441	99.3	0.7	99.2	0.5	99.6	0.7
Molders, metal	36,382	99.8	0.7	99.8	0.8	99.8	0.9

Notes: This table shows summary statistics for the top 20 occupations with some variation in Social Security coverage year in the 1930 Census (defined as occupations with a strictly positive share of workers covered in 1935 and a strictly positive share of workers covered in 1950 or later). Population consists of men born in 1875–1888 whose current employment was covered in 1935 or 1950 (excluding the self-employed). Columns 4 and 5 restrict the sample to men who appear in our Family Tree sample. Columns 6 and 7 restrict the sample to men who appear in our claims linked sample. Observations in columns 4–7 weighted using inverse propensity score weights that re-weight the sample to match the population of interest in 1930.

Table A.3: Summary of physical demands and working conditions by occupational categories in 1966 supplement to 1965 Dictionary of Occupational Titles

		Physical demands										Working conditions								
		1					2	3	4	5	6	1			2	3	4	5	6	7
Occupational categories	*N*	S	L	M	H	V						I	O	B						
Professional, technical & managerial	1,757	46.7	47.9	4.7	0.4	0.2	4.4	3.9	52.9	67	54.4	81.4	2.8	15.7	0.1	0.3	1.5	4.1	5.5	3.2
Clerical & sales	977	38.8	50.2	8.5	2.1	0.4	2.3	7	72.2	48.8	29.7	84.6	2.8	12.5	0.5	0.2	0.4	2.1	1	1.3
Service	532	4.3	60.9	25.9	7.5	1.3	8.6	19.5	80.3	48.9	36.1	77.8	6.8	15.4	3.2	13.2	16.5	10	16.4	9
Farming, fishery, forestry & related	302	1	36.4	35.4	22.8	4.3	33.4	46.7	84.8	18.9	37.1	17.9	64.6	17.5	0.3	0.3	24.2	9.9	26.5	17.2
Processing	3,059	1.5	45.2	38.1	13.4	1.8	11.4	20.5	95.6	10.3	26.4	92.3	2.6	5	1.3	13.8	14.6	18.6	19.3	25.2
Machine trades	2,206	4.3	44.1	40	10.7	1	5.5	21.7	96.9	11.7	42.5	94.4	1.6	3.9	0.1	3.8	2.8	34.5	16.8	9
Bench work	2,824	16.4	58.5	20.9	3.8	0.4	1.1	8.6	97.8	6.3	54	98.4	0.3	1.2	0	1.2	2.4	9	5.6	5.6
Structural work	1,012	0.9	26.8	47.1	22.8	2.4	38.7	65.9	95.2	20.8	42.8	43	33.9	23.1	0.1	3.3	7.6	28.1	41.8	17.3
Miscellaneous	1,146	10.6	44.2	29.9	11.9	3.3	15.4	24.9	89.2	22.3	40.6	59.9	22.4	17.6	0.9	1.6	7.9	21.7	18.4	11.2
Total	13,815	14.2	47.4	28	9.1	1.3	9.5	19.4	87.8	23.1	41.4	83.5	7.5	9	0.6	4.8	6.8	16.6	14.7	11.6

Notes: This table shows the total number of job titles and the share of job titles with specific physical demands or working conditions by 1-digit occupational category (see Appendix Figure A.4 for definitions).

Table A.4: Occupations in top quartile based on physical demands and working conditions index

Occupation	Number (1)	Share covered in 1935 (2)	Share physically-demanding (3)	Share exposed to hazards (4)	Share exposed to toxic conditions (5)	Share (3) or (4) or (5) (6)
Roofers and slaters	3,951	99.6	0.0	97.9	0.4	98.2
Welders and flame cutters	6,999	99.4	0.0	97.0	74.9	97.1
Tinsmiths, coppersmiths, and sheet metal workers	13,840	99.1	0.0	90.4	5.2	90.7
Painters, construction and maintenance	92,578	97.2	0.2	82.8	50.4	87.1
Electricians	42,196	95.4	0.0	85.1	1.9	85.6
Boilermakers	8,066	99.3	0.0	83.5	3.6	83.5
Plasterers	15,572	98.7	0.0	77.1	21.5	77.9
Structural metal workers	10,019	99.9	0.1	70.2	11.9	70.2
Furnacemen, smeltermen and pourers	4,454	99.1	0.3	67.1	52.2	68.6
Plumbers and pipe fitters	48,628	96.8	0.1	67.7	14.1	68.6
Cooks, except private household	42,331	95.8	1.0	51.5	28.8	66.1
Carpenters	245,313	98.5	5.4	65.4	1.4	65.9
Lumbermen, raftsmen, and woodchoppers	34,914	99.8	5.8	63.7	3.4	64.5
Stone cutters and stone carvers	6,249	99.3	6.1	31.3	53.1	62.4
Painters, except construction or maintenance	13,137	99.7	0.1	11.3	59.3	62.3
Stationary firemen	34,659	92.9	10.8	57.4	33.0	60.4
Brickmasons, stonemasons, and tile setters	43,349	99.0	6.3	47.7	19.8	55.2
Molders, metal	36,382	99.8	0.0	17.3	42.2	54.2
Teamsters	21,477	92.3	0.0	53.9	0.0	53.9
Cement and concrete finishers	4,611	98.9	1.4	49.9	3.8	51.6
Inspectors (n.e.c.)	9,046	99.8	0.0	41.2	24.0	50.3
Guards, watchmen, and doorkeepers	35,093	87.7	0.1	50.0	13.0	50.2
Loom fixers	6,207	100.0	0.0	0.0	50.0	50.0
Motion picture projectionists	1,747	99.9	0.0	50.0	50.0	50.0
Cranemen, derrickmen, and hoistmen	8,329	99.3	0.0	33.8	35.3	49.0
Sawyers	8,293	99.8	2.2	40.0	10.8	47.5
Stationary engineers	73,678	92.5	0.1	40.0	14.2	45.9
Excavating, grading, and road machinery operators	2,108	96.9	0.8	23.2	29.3	45.8
Attendants, auto service and parking	6,316	99.3	0.1	2.3	43.0	43.9
Millwrights	14,434	99.8	0.0	43.7	10.9	43.7
Meat cutters, except slaughter and packing house	27,530	99.1	0.0	42.8	21.4	43.1
Linemen and servicemen, telegraph, telephone, and power	15,083	98.3	4.5	38.3	4.4	42.8
Foresters and conservationists	1,201	87.0	11.3	39.4	27.0	39.4
Mechanics and repairmen (n.e.c.)	35,930	97.2	1.2	30.6	11.3	35.7
Laborers (n.e.c.)	795,300	89.3	7.9	27.0	13.4	34.7
Truck and tractor drivers	82,226	95.5	8.4	22.5	4.5	34.3
Craftsmen and kindred workers (n.e.c.)	11,389	99.7	0.4	32.6	6.1	33.4
Power station operators	1,119	97.6	0.0	29.3	4.9	33.0

Notes: This table shows summary statistics for occupations in the top quartile based on the share of associated job titles that are "very physically demanding" or in which workers are exposed to "hazards" or "toxic" working conditions (see Section 5.2 for details). Sample restricted to occupations with 1,000+ workers in our population of interest in the 1930 Census and some variation in Social Security coverage (share of workers covered in 1935 strictly between 0 and 1).