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The Mortality Effects of Retirement: Evidence from Social Security Eligibility at Age 62*

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Abstract

Social Security eligibility begins at 62, and approximately 40 percent of Americans begin claiming benefits within several months of their 62nd birthday. We use a regression discontinuity design to examine whether mortality changes discontinuously at this threshold. Using mortality data that covers the entire U.S. population and includes exact dates of birth and death, we document a robust two percent increase in male mortality at age 62. Discontinuous changes in mortality do not occur at nearby ages, suggesting that the increase at age 62 results from Social Security eligibility and the associated changes in lifestyle that occur at that age.

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1. Introduction

There is enormous interest in the effect of retirement on health, especially given the aging of the population and the reforms to retirement policies underway in the United States and other developed countries. However, the interdependence between health outcomes and retirement status, which commonly leads individuals in poorer health to retire earlier, means that associations between retirement and health do not necessarily represent causal relationships.

There is a growing literature on how retirement affects health.[1] Some researchers use longstanding pension rules in the U.S. and Europe that affect the likelihood of retirement to identify the relationship (e.g., Charles, 2004; Bound and Waidmann, 2007; Neuman, 2008; Coe and Zamarro, 2011; Behncke, 2012), while other researchers use the offer of early retirement incentives (e.g., Coe and Lindeboom, 2008; Hernaes et al., 2013) or other changes in retirement policy.[2] Despite the wide range of settings and variety of empirical strategies, these studies have failed to produce clear and consistent findings as to how retirement affects health. A number of studies find that subjective health and wellbeing measures are improved by retirement. However, many of these same studies find retirement has no effect on objective health outcomes (e.g., Neuman, 2008; Coe and Zamarro, 2011; Insler, 2014). Recent studies in Europe examining the mortality effects of retirement have produced mixed findings.[3] Many studies lack sufficient statistical power to generate precise estimates, which also makes it difficult to identify the mechanisms underpinning particular findings.

We contribute to this literature by examining how mortality changes in the United States at age 62, when most individuals first become eligible for the Retirement and Survivors component of Social Security ("Social Security"). While individuals receive a higher rate of benefits if they delay claiming until later ages, many begin to claim Social Security upon turning age 62. Figure 1 shows the monthly rate of new Social Security claims constructed using a one percent extract of Social Security administrative data for cohorts born between 1921 and 1948.[4]

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[1] There is a related literature on how health affects retirement and the claiming of pensions. For example, see Rust and Phelan (1997); Bound et al. (1999); McGarry (2004); French (2005); Scholz and Seshardi (2012).

[2] Additional approaches include the use of individual fixed effects (Dave, Rashad and Spasojevic, 2008) and extended unemployment benefits for blue-collar workers that allow early retirement (Kuhn et al., 2010).

[3] Blake and Garrouste (2013) find that retirement decreases mortality among seniors in France, while Bloeman, Hochguertel and Zweerink (2013) find a similar result when examining early retirees from the public sector in the Netherlands. Kuhn, Wuellrich and Zweimueller (2010) find mortality increases among early-retiring blue collar workers in Austria. Hernaes et al. (2013) find no mortality effects from a retirement reform in Norway.

[4] This includes new claims for Social Security Disability Insurance. Population data is from the Current Population Survey. We explain more about these data in Section 2.

The spike in new Social Security claims at age 62 is striking: 31 percent first claim in the month after turning age 62, while a further five percent claim in the adjacent months.

There is also a discontinuous change in the age-mortality relationship at age 62. We identify this relationship using a regression discontinuity design and restricted-use versions of the National Center for Health Statistics' Multiple Cause of Death (MCOD) data. The MCOD data are compiled from death certificates and include exact dates of birth and death. We estimate that the aggregate mortality of cohorts born between 1921 and 1948 increases by approximately 1.5 percent at age 62. This is largely driven by a two percent increase in male mortality that is statistically significant and robust to a wide variety of specification and bandwidth choices. While there is a one percent increase in female mortality, it is sensitive to these choices.

We find that the size of the increase in male mortality differs by decedents' demographic characteristics. There are differences by marital status, with much larger estimates for unmarried males (3.8-4.2 percent, $p<0.01$) than married males (0.8-1.3 percent, $p\geq0.05$). There are also differences by educational attainment, with a large increase for males who did not complete high school (2.8-3.0 percent, $p<0.05$) compared to smaller, statistically insignificant increases for graduates of high school and college. There are also differences by the place and cause of death. In terms of place of death, the increase in male mortality is largest outside of hospitals, nursing homes and other institutions (i.e., at home and public places) (3.2-3.4 percent, $p<0.01$). This compares to a smaller, statistically insignificant increase in male mortality in hospitals and no change in nursing homes/institutions. The cause-of-death categories with statistically significant increases in male mortality are chronic obstructive pulmonary disease (COPD) (5.0-7.0 percent, $p<0.05$), lung cancer (5.1-5.3 percent, $p<0.01$) and external causes (3.1-4.0 percent, $p<0.05$).

We examine whether the increase in mortality is related to the resources available upon Social Security eligibility. We create subsamples based on differences in cohorts' "Full Retirement Age" (FRA), which changed from 65 to 66 years over this period and decreased Social Security benefits at age 62 by 6.25 percent.[5] A two percent increase in male mortality is present for both earlier cohorts and later cohorts, suggesting that the increase in male mortality at 62 is a persistent phenomenon not affected by recent changes to the benefits paid at age 62.

[5] When the FRA was 65, individuals claiming Social Security at age 62 received 80 percent of their Primary Insurance Amount. Individuals with a FRA of 66 receive 75 percent of their Primary Insurance Amount if they claim Social Security at age 62.

We use data from the Social Security Administration (SSA) and the Health and Retirement Study (HRS) to examine what these results suggest about why mortality increases at age 62. We estimate that the discontinuous change in male Social Security claiming rates at age 62 is 31 percent, which suggests that mortality increases by approximately six percent at age 62 if attributed to those who begin claiming. However, females have a similar increase in Social Security claiming at age 62, albeit with lower average payments, and male claiming rates do not differ substantially by marital status or educational attainment. This suggests that it might not be Social Security claiming per se that accounts for the increase in mortality. There are many other changes at age 62, including a large increase in the fraction retired, a large decrease in the fraction doing paid work and an increase in the fraction without health insurance. Changes in retirement and paid work do generally differ by sex and demographic characteristics in ways consistent with the mortality estimates, suggesting that it may be these changes that account for the increase in mortality at age 62.

Our estimates apply to age 62 Social Security claimants, who differ from later claimants and non-claimants in terms of their demographic, socioeconomic and health characteristics (Li, Hurd and Loughran, 2008). However, they represent a sizeable fraction of the U.S. population. It is difficult to know for how long after age 62 the elevated mortality persists, and also the consequences of delaying the age at which individuals become eligible for Social Security.

The rest of the paper is organized as follows. In Section 2, we provide background about Social Security and claiming at age 62. In Section 3, we describe the data and the sample and then, in Section 4, analyze the change in mortality at age 62. In Section 5, we examine the changes in Social Security claiming and other activities at age 62. We conclude in Section 6.

2. Age 62 and Social Security Eligibility

Social Security Retirement Insurance is available to older workers who have ten or more years of Social Security-taxable employment.[6] While there was a single retirement age of 65 when Social Security was first established, the ability to claim Social Security at age 62 was established for women in 1956 and for men in 1961. Age 65 became the FRA (or the "Normal Retirement Age"), the age at which workers would receive the full "Primary Insurance Amount"

[6] Workers need 40 credits that used to be based on "quarters of coverage" but now depend on annual earnings. In 2015, workers need $4,880 in covered earnings to receive the maximum four credits for the year. For more details about this and other SSA policies discussed in this section, see SSA (2012).

(PIA) due to them under the Social Security earnings-payment formula. Benefits were reduced by 5/9 of one percent for each month before the FRA that an individual began claiming Social Security, so that workers claiming at age 62 received 80 percent of their PIA. Cohorts born after 1942 have a FRA of 66 and age 62 claimants now only receive 75 percent of their PIA.[7]

Social Security Retirement and Survivors Insurance is also available to qualifying workers' dependents and survivors. These are most commonly a worker's spouse, divorced spouse, or widow(er).[8] Dependents' base level of benefits is up to 50 percent of the worker's PIA. Spouses (or former spouses) can receive Social Security benefits upon turning 62, but there is a (slightly larger) penalty for claiming early so that they receive 70-75 percent of their base benefits if they claim at age 62 (depending on their FRA). Widows(er)s can receive benefits from age 60; again, there are reductions for claiming earlier than the FRA. Payments to Social Security beneficiaries cease the month after death.[9]

Individuals can file an application for Social Security benefits two or three months before turning age 62 and become eligible for Social Security either in the month they turn 62 or the following month.[10] As described in the introduction and shown in Figure 1, many people claim Social Security at the first possible opportunity, despite the penalty to claiming early.[11] We can use these tabulations of a one percent extract of SSA's Master Beneficiary Record to examine claiming rates by sex. This is shown in Appendix Figure A1. The claiming rates are reasonably similar for men and women, although women are more likely to claim as a dependent, either as a spouse, former spouse or widow.[12] Age 62 claimants who are women generally receive a lower

[7] Starting with the 1938 cohort, the FRA increased in two month increments, so the 1938 cohort has a FRA of 65 and 2 months, the 1939 cohort has a FRA of 65 and 4 months, etc. There is now an early claiming penalty of 5/9 of one percent per month for the first 36 months of early claiming and 5/12 of one percent for each additional month. There are also "Delayed Retirement Credits" for delaying Social Security claiming up to age 70.

[8] Other dependents include children under 18, adult children who became disabled before 22 and parents who are reliant on the worker for income support. These groups are more commonly dependents for the younger recipients of Social Security Disability Insurance beneficiaries than Retirement and Survivors Insurance beneficiaries.

[9] Family members, funeral homes and government agencies report deaths, so there are limited opportunities to delay reporting. Dependents can receive a lump-sum death benefit of $255. Dependents can receive a higher rate of benefits upon the death of a worker or worker beneficiary.

[10] Individuals born on the 1st or 2nd of the month become eligible in the month they turn 62, while individuals born later in the month become eligible in the month after they turn 62. The first Social Security payments generally arrive the month following initial eligibility. Payments occur monthly; traditionally on the 3rd, although beneficiaries eligible after May 1997 are paid on a Wednesdays that depends on their date of birth. See Olson (1999) for details about the timing of claiming at age 62 and Evans and Moore (2011) for details about the payment schedule.

[11] Note that these rates are based on the overall U.S. population, so rates among workers and their dependents with sufficient covered employment are even higher.

[12] Song and Manchester (2008) and publicly-available SSA statistics show similar spikes in claiming at age 62 for males and females (e.g., Table 6.A4 in SSA, 2012).

monthly Social Security benefit than men; in 2010, the average payments for women and men were $872 and $1,168 per month, respectively (SSA, 2012, Table 6.B1). Compared to individuals who claim Social Security at later ages, both male and female age 62 claimants are less educated, in worse health, have had lower earnings and been more likely to work in physically demanding jobs (Li, Hurd and Loughran, 2008).

Social Security beneficiaries are not required to stop working, though they are subject to a "Retirement Earnings Test" from age 62 until they reach the FRA. Beneficiaries have one dollar of benefits withheld for every two dollars of earnings above a threshold, which is currently $15,720 per annum.[13] Even though beneficiaries later receive benefits that are withheld, many beneficiaries appear to view the Retirement Earnings Test as a tax and reduce their labor force participation as a result (e.g., Friedberg, 2000; Engelhardt and Kumar, 2014; Gelber, Jones and Sacks, 2014). Most individuals do substantially decrease their employment after claiming Social Security at age 62 and also report large changes in retirement status, income and health insurance (Rust and Phelan, 1997; Gustman and Steinmeier, 2005; Li, Hurd and Loughran, 2008).

Any change in aggregate mortality at age 62 can be attributed to Social Security only if there no other policies or programs affecting health specifically or discontinuously at age 62. To our knowledge, there are no other federal programs with eligibility rules that change discontinuously at age 62. While there may be private pension systems that encourage retirement at age 62, private sector employers have largely switched to defined contribution pensions that provide little incentive to retire at a specific age. The defined benefit programs used in the public sector typically eligible for a full pension is usually much younger than 62. We are not aware of non-retirement policies where eligibility is based on turning age 62. Although it is difficult to rule out the possibility of state-level policies have specific effects at age 62, we estimate the effects for different Census regions to ensure that any results are not driven by the effects of policies in a specific state or region.

3. Mortality Data and Sample Characteristics

The mortality data comes from restricted-use versions of the National Center for Health Statistics' Multiple Cause of Death (MCOD) files from 1979 to 2012. The files include

[13] There are different rules for the year that a beneficiary reaches the FRA. There have been several changes to the Retirement Earnings Test, which have generally affected how it is applied at the FRA and above.

decedents' dates of birth and death; demographic information, including age, sex, race, marital status and educational attainment (since 1989); and information on their cause and place of death. Reported sex, race and place of residence closely match survey data (Sorlie, Rogot and Johnson, 1992), while educational attainment is higher than reported in other data (Sorlie and Johnson, 1996).[14] Underlying cause of death is coded using the 9th version of the International Classification of Disease until 1998 and the 10th version thereafter. We create cause-of-death categories that are consistent across both versions.[15]

Given that age of death is our assignment variable, it is important to emphasize that the date of birth and date of death are reported to the exact day. This enables us to precisely measure age at death. There is no odd heaping in the distributions of these dates, which suggests that any patterns in death counts should not reflect reporting differences related to the data collection.[16] Such heaping in data is important to consider before implementing a RD design (Barreca, Lindo and Waddell, 2015).

In order to have sufficient bandwidth and the opportunity to conduct placebo tests at nearby ages, we focus on individuals born between 1921 and 1948 (i.e., turning 62 between 1983 and 2010).[17] In Table 1, we show the number and composition of deaths in this group at ages 61 and 62 for all decedents, and separately for males and females. There are 65,489 deaths per month at age 61 and 70,727 deaths per month at age 62. Males account for approximately 60 percent of the deaths at both ages, well above their proportion of the population at these ages. The other demographic characteristics are broadly in line with general population characteristics, although it has been well documented that there is higher mortality at relatively young ages among those with characteristics related to low socioeconomic status, such as not completing high school and not being married (e.g., Hu and Goldman, 1990; Lantz et al., 1998).

We use Social Security administrative data to measure changes in claiming at age 62. Information is drawn from a one percent extract of the Master Beneficiary Record, which contains the program data used to manage both Disability Insurance and Retirement and

[14] This study used educational attainment reported in 1989, the first year that it was available in the MCOD files.
[15] Details of these categories are provided in the appendix.
[16] This is in contrast to another major source of mortality data, the Social Security Master Death File, which has unusual peaks at the 1st and 15th of the month that likely partly reflect the way the data is collected (twice a month from hospitals, funeral homes, etc.).
[17] Our sample includes only one cohort (born 1921) affected by the Social Security notch, which has been shown to have its own mortality effects (Snyder and Evans, 2006).

Survivors Insurance. We have these data for the same 1921-1948 birth cohorts. Population data is drawn from the Current Population Survey.

4. The Change in Mortality at Age 62

4.1 Graphical Evidence

We present the number of deaths in relation to age 62 in Figure 2. We include monthly counts for 12 months on either side of age 62. There is a large increase in mortality precisely at age 62, with 1,580 (2.4 percent) more deaths in the month after turning 62 than the month before. By comparison, the number of deaths increases by an average of 316 deaths (0.5 percent) per month in the eleven months prior to turning age 62. In Figure 3, we show that this increase in mortality occurs for both males (Panel A) and females (Panel B), although there is a clearer and larger increase in mortality for males than for females.

It is possible that we are identifying a change that is related to having a birthday or generally becoming one year older. To check this, in Appendix Figure A2 we present the same monthly mortality counts in relation to ages 61 and 63. We are not aware of any important age-based policy rules at these ages. There is no obvious change in mortality at either of these ages, either for the full sample or for samples of males and females.

4.2 Regression Estimates

To estimate the mortality effects of the availability of Social Security at age 62, we use a RD design where we compare the number of deaths of individuals aged just older than age 62 to individuals slightly younger. RD designs and age-based eligibility thresholds have been used to identify, for example, the effect of Medicare eligibility at age 65 on mortality (Card, Dobkin and Maestas, 2009) and the effect of students losing access to parental health insurance at age 23 on emergency department visits (Anderson, Dobkin and Gross, 2014). A comparison of deaths above and below age 62 should generate causal estimates of the local average treatment effect of Social Security eligibility on mortality as long as other factors affecting mortality do not change discontinuously at age 62.

We implement several global parametric and local nonparametric RD specifications that allow us to assess the stability of the results under different assumptions about the data generating process. The global parametric regressions contain polynomials that control for the underlying age-mortality relationship and a dummy variable to estimate the change in mortality at age 62. In our setting, the basic form of the regression is:

$$\log(Mortality_a) = f(a) + Post62_a\,\beta + \varepsilon_a \quad (1)$$

The dependent variable is the natural log of mortality counts at age at death a, which allows us to measure the change in mortality in percentage terms. We use monthly counts, which is the level of aggregation suggested by RD tests of excess smoothing.[18] The function $f(a)$ represents the polynomials used to control for the relationship between age and mortality. We allow this relationship to vary on either side of the discontinuity and estimate it using quadratic, cubic and quartic polynomials in separate specifications.[19] The dummy variable $Post62_a$ is equal to one above age 62 and the coefficient β gives the percentage change in mortality at age 62 (our coefficient of interest). The error term is ε_a. We use Huber-White robust standard errors and allow for an arbitrary correlation in errors in relation to age.

In the first column of Table 1, we present the estimated change in mortality at age 62. In this specification, we use a bandwidth of 12 months. The point estimates are 1.35 percent (quadratic), 1.97 percent (cubic) and 1.93 percent (quartic). All are statistically significant at the one percent level. The cubic specification minimizes the finite-sample (corrected) Akaike Information Criterion (AICc).

We also use local nonparametric specifications to relax the functional form assumptions and place more weight the observations near the cutoff. We use the local linear and local quadratic specifications of Calonico, Cattaneo and Titiunik (CCT) (2014a, 2014b), who develop a "data-driven" bandwidth selection procedure based on minimizing the mean square error of the point estimate. We use their bandwidth selection procedures starting with 24 months of data on either side of the discontinuity.[20] We report estimates that use a triangular kernel, their

[18] Lee and Lemieux (2010) suggest two tests for the level of aggregating outcome data in RD settings. The first assesses whether using narrower bins provides a better fit to the data. It is implemented by comparing the R-squared from a regression with dummy variables for each bin of width *w* to the R-squared from a regression with dummy variables for each bin of width *w*/2. The bin size *w* is decreased until the resulting F statistic is not statistically significant. The second test is based on the idea that a bin width is too wide if, within each bin, there is a systematic relationship between the outcome variable and the assignment variable. It is implemented by interacting each bin dummy variable with the assignment variable, regressing the outcome variable on the set of bin indicators as well as these interaction terms, and testing the joint significance of the interaction terms. The bin size *w* is decreased until the F-statistic on the joint test is not statistically significant. In this setting, monthly counts satisfied both tests. That said, in addition to monthly counts, we use other levels of aggregation in our regression analysis.

[19] We use a variety of polynomial lengths as there is a debate about whether or not higher-order polynomials should be used in RD designs (see Lee and Lemeieux, 2010; Card et al., 2014; Gelman and Imbens, 2014).

[20] Social Security rules and other policies that take effect at age 60 prevent using a larger starting bandwidth.

procedures to correct for bias due to bandwidth choice, and plug-in residuals for standard error estimation, which are equivalent to Huber-White robust standard errors in the parametric context.

The results from these regressions are also presented in the first column of Table 2, together with the bandwidths used. The point estimates and statistical significance are similar to the parametrically-generated results, with point estimates of 1.42 percent (local linear) and 1.93 percent (local quadratic) that are statistically significant at the one percent level. Across these five specifications, and in accord with the visual evidence presented previously, we find a statistically significant increase in mortality at age 62 of one to two percent.

We verify that these estimates are robust to additional regression controls and other choices (Appendix Table A1). First, we show that the global parametric estimates are similar when year-of-birth or month-of-death fixed effects are included. Second, we show that all five estimates are similar when monthly mortality counts are replaced by daily or weekly counts. Third, we present estimates where the cutoff is defined by the precise Social Security eligibility rules. These rules allow individuals born on the 1^{st} or the 2^{nd} of the month to claim Social Security in the month they turn 62, while individuals born later in the month become eligible the following month (Olson, 1999). The estimates are qualitatively similar, except that the global quartic and local quadratic estimates increase by approximately one percent and the standard errors increase so that the global quadratic estimate is no longer statistically significant at conventional levels. These suggest that our estimates are robust to important modeling choices.

In Table 2, we also present estimates of the change in mortality at age 62 separately for males (in column 1) and females (in column 2) using the same three global parametric and two local nonparametric specifications. For males, the estimated increase in mortality at age 62 is between 1.85 and 2.43 percent and is always statistically significant at the one percent level. For females, the estimated increase in mortality at age 62 is between 0.58 and 1.38 percent. Only the quadratic estimate of 0.58 percent not statistically different from zero. The other four estimates are statistically significant at the one percent level, including the quartic polynomial that minimizes the finite-sample AIC. These results are in accord with the visual evidence that there is a larger and clearer change in mortality for males than for females.

We next assess the robustness of the estimates to different bandwidths by using bandwidths of between six and 24 months. Having established that the results are generally similar with and without higher-order terms, in this and following exercises we focus on the local linear and global quadratic specifications. Appendix Figure A3 shows the estimates and 95 percent confidence intervals from both of these specifications for the full sample, for males and for females. For the full sample and for males, the local linear and global quadratic estimates are between 1.5 and 2.4 percent and statistically significant for all bandwidth values. The results for females are more sensitive to the choice of bandwidth. The estimated change in mortality declines as the bandwidth is increased and loses statistical significance at the five percent level at 11 months in the local linear regression and 12 months in the global quadratic regression. At larger bandwidths, the estimates for females are small and imprecisely estimated.

Another robustness exercise is to assess whether placing a discontinuity at age 62 matches the data better than "placebo" discontinuities at nearby locations.[21] We do this by comparing the model fit, as measured by r-squared, to specifications that have a placebo discontinuity at monthly increments throughout the data window.[22] To conduct this exercise, we use the global quadratic specification and 12 months of data on each side of age 62. In Appendix Figure A4, we report the r-squared at each location for the full sample, males and females. For the full sample and for males, the r-squared value is clearly maximized at the actual age 62 threshold, indicating that the change in mortality at age 62 is not spurious. For females, however, the r-squared at age 62 is the fourth-largest of those throughout the 24-month window. This indicates that the change in mortality at age 62 is not unique for females.

The estimates for increase in the mortality of males at age 62 are robust across many different RD approaches. On the other hand, the results for female are sensitive to bandwidth and specification choice. Henceforth, we therefore focus on the increase in male mortality.

We are estimating a local average treatment effect. The structure of the RD design makes it inherently difficult to determine how "local" this effect is, but we can add dummy variables to the global quadratic regression close to age 62 to understand the extent to which the results are driven by observations close to age 62. If we add individual dummy variables to the first three months after age 62, the estimated coefficient (standard error) is 0.0386 (0.0114). If we also add

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[21] We also estimate placebo regressions at ages 61 and 63 in Appendix Table A2. In line with the visual evidence shown in Appendix Figure A1, the estimated discontinuity in mortality at these ages is small and statistically insignificant at conventional levels. While there is an estimated reduction in male mortality at age 61 of around two-thirds of one percent that is statistically significant at the five percent level across both the local linear and global quadratic specifications, this relationship is not robust to including higher-order polynomials. The global cubic and quartic regressions produce coefficients (standard errors) of -0.0037 (0.0039) and -0.0082 (0.0048), respectively.

[22] Kane (2003) proposed this test. He reported the log likelihood throughout the data window, which is an equivalent measure of model fit for this regression specification.

individual dummy variables to the first three months before age 62, the coefficient (standard error) is 0.0247 (0.0145). This second estimate is equivalent to a so-called "donut hole" RD where three months on either side of the age 62 discontinuity are dropped. If we do this instead of using dummy variables, the coefficient (standard error) is 0.0247 (0.0126).

4.3 Heterogeneity of Mortality Effects

The results suggest there a substantial number of "excess" deaths for males after age 62. For example, the estimated increase from the local linear regression of 2.15 percent translates into 17,864 additional male deaths at age 62 – or an average of 638 extra deaths per birth cohort – than without the discontinuous increase at age 62. We now analyze subgroups in order to understand what is driving this aggregate pattern. We continue to focus on estimates produced from a local linear regression and Calonico, Cattaneo and Titiunik's (2014) "data driven" bandwidth and a global quadratic regression using a bandwidth of 12 months.

Note that any difference across these groups could reflect differences in Social Security claiming; differences in the change in labor force participation; or differences in how these changes affect mortality through, for example, differences in underlying health or health behaviors. We will discuss the heterogeneity in the mortality effects in this section, and then seek to understand the likely cause of observed differences in the next section.

Results for a range of demographic and mortality characteristics are presented in Table 3. In addition to the local linear and global quadratic estimates, we also report the fraction of the total deaths at ages 61 and 62 for each group. Marital status is presented first. Given that 65 percent of the decedents were married, we initially divide the sample into married and non-married males. For married males, the local linear estimate is 1.30 percent ($p<0.05$) and the global quadratic estimate is an imprecise 0.81 percent ($p\geq 0.05$). In contrast, the non-married estimates are substantially larger, with a local linear estimate of 4.15 percent ($p<0.01$) and a global quadratic estimate of 3.77 percent ($p<0.01$). We examine the non-married group in further detail by looking separately at single, divorced and widowed males. The point estimates for single males are the largest (5.14-5.58 percent), followed by divorced males (3.05-3.37 percent) and widowed males (2.62-3.30 percent). Estimates for the single and divorced groups are (at least) statistically significant at the five percent level, while they are not for widowed males, who only account for 6.2 percent of the overall sample.

We next present results based on differences in educational attainment. Note that educational attainment has only been available in the MCOD file since 1989, so that the sample is slightly smaller. The increase in mortality at age 62 is largest for males who did not complete high school, with a local linear estimate of 3.03 percent ($p<0.01$) and a global quadratic estimate of 2.75 percent ($p<0.05$). The estimates for those who completed high school and those who completed college are smaller and not statistically significant at any level, although the standard errors for college males are large and 95 percent confidence intervals do not rule out large increases in mortality at age 62.

Race is defined in terms of white and non-white males, as non-white males only account for 16.5 percent of the overall sample. For white males, the estimates are large (2.09 to 2.37 percent) and statistically significant ($p<0.01$). For non-white males, the estimates are smaller (0.68 to 0.78 percent) and not statistically significant, although the confidence intervals are again wide and allow for a similar increase at age 62 as for white males.

We now turn to the place and cause of death. The place-of-death subgroups consist of deaths outside of hospitals and institutions (including those classified as dead on arrival), deaths in hospital and deaths in nursing homes and other institutions. The largest increase at age 62 occurs for deaths outside of hospitals or institutions, with estimates of 3.23 to 3.39 percent, $p<0.01$ (for both specifications). Deaths in hospitals are estimated to increase by 1.24 to 1.30 percent, $p<0.10$ (for both specifications). Deaths in nursing homes and other institutions are estimated to change by a statistically insignificant -0.22 to -0.79 percent.

We examine causes of death that are consistent across Versions 9 and 10 of the International Classification of Disease. We use four large groupings: heart and lung conditions (heart disease, stroke and chronic obstructive pulmonary disease (COPD)), cancers, external causes and all other causes. The estimates for heart and lung conditions are 1.35 to 2.50 percent, with the larger (local linear) estimate the only statistically significant one ($p<0.05$). Larger, statistically significant increases are present for both cancers (2.62 to 2.63 percent, $p<0.01$ for both specifications) and external causes (3.14 to 3.99 percent, $p<0.01$ and $p<0.05$ respectively), but not for all other causes (1.09 to 1.13 percent, $p\geq0.05$). We further examine narrower classifications of heart/lung conditions and cancers.[23] When heart/lung conditions are separated into heart attacks, COPD and other heart/lung causes (other heart disease, strokes), the increase

[23] This is not possible for external causes, as they only account for five percent of all deaths.

for COPD (4.96 to 6.96 percent, $p<0.05$ and $p<0.01$ respectively) is striking. In contrast, the point estimates for heart attacks and other heart/lung causes are smaller and not statistically significant. When we split cancers into lung cancer and other cancers, it becomes apparent that the increase in cancers is due to a large increase in lung cancer deaths at age 62 (5.10 to 5.31 percent, $p<0.01$ for both). In contrast, the estimated increase in other cancers is 0.99 to 1.06 percent, $p\geq0.05$.

Finally, we consider whether the change in mortality at age 62 is specific to particular cohorts or periods of time.[24] In Table 4, we present estimates for males born between 1921 and 1937 (turning 62 between 1983 and 1999) and 1938 and 1948 (turning 62 between 2000 and 2010). The full retirement age (FRA) is 65 for the earlier cohorts. The later cohorts have a FRA above 65: the 1938-1942 cohorts have a FRA of between 65 and 66, while the 1943-1948 cohorts have a FRA of 66.[25] Recall that the change in the FRA increased the reduction in Social Security payments associated with claiming at age 62 from 20 to 25 percent. There is evidence that it changed claiming behavior (e.g., Blau and Goodstein, 2010; Song and Manchester, 2008; Mastrobuoni, 2009). However, there is no discernible difference in the increase in mortality at age 62 across these cohorts. The estimates for the 1921-1937 cohorts are 1.48-1.70 percent ($p\geq0.05$ and $p<0.05$, respectively), which are close to the estimates for the 1938-1948 cohorts of 1.79-2.44 percent ($p<0.01$ and $p<0.05$, respectively). We also divide the recent cohorts into those born in 1938-1942 and 1943-1948. The estimates are similar across these two groups. These results suggest that the increase in male mortality at age 62 is a persistent phenomenon, and there is no evidence that is affected by the changes associated with increases in the FRA.

In summary, there is substantial heterogeneity in the increase in mortality at age 62. In terms of demographic characteristics, the largest increases occur among non-married males, especially single ones, and males who did not complete high school. The increase in deaths occurs predominantly outside of hospitals and institutions, although there is also a smaller increase within hospitals. In terms of underlying causes of death, the largest (and only statistically significant) increases occur for COPD, lung cancer and external causes. As we show in the next section, there are changes in behavior at age 62 consistent with these differences.

[24] Focusing on age 62 means that differences across cohorts are indistinguishable from differences over time.
[25] The FRA was increased in two month increments, so that the 1938 cohort has a FRA of 65 and 2 months, the 1939 cohort has a FRA of 65 and 4 months, and so on.

5. Social Security Claiming and Age 62

The estimates clearly show an increase in male mortality at age 62. Not everyone claims Social Security at age 62, so now we use Social Security administrative data to get some idea of the scale of the mortality effects under the assumption that they are only due to the change in Social Security claiming at that age. We also examine whether differences by sex and other demographic characteristics could result from Social Security claiming differences, other changes that are proximate to beginning to claim Social Security, or differences in underlying health. To do that, we use the administrative data and also information from the HRS.

In Appendix Figure A4, we have already shown that males and females have similar spikes in new Social Security claims at age 62. We apply the same RD specifications as used in Section 4 to estimate the discontinuity in claiming at age 62 for males and females. To do this, we convert the monthly claiming rates into cumulative Social Security rates, as presented in Appendix Figure A5. Using these rates as the dependent variable, we estimate global quadratic, cubic and quartic regressions with a bandwidth of 12 months, and local linear and quadratic regressions using CCT-calculated bandwidths. The estimates are presented in Appendix Table A3. For males, the point estimates suggest that Social Security claiming increases by 30.1 to 31.1 percentage points ($p<0.01$). For females, the point estimates are 32.2 to 33.5 percentage points ($p<0.01$). If we scaled the increase in male mortality by – and therefore attribute it to – the change in claiming, this suggests that male mortality increases by approximately six percent at age 62 among new Social Security claimants.

However, there is no difference in claiming rates by sex, suggesting that there be other reasons for the age 62 increase being concentrated among males. To explore this aspect further, we use the HRS to see if the demographic differences in males' claiming patterns account for the heterogeneity in the mortality effects. We calculate the population-weighted fraction of male HRS respondents who first claim at age 62 by marital status, educational attainment and race. These are observed in the last interview before turning 62 (i.e., at 60 or 61), and presented in Appendix Table A4. While the rates are slightly lower than found in the administrative data, suggesting there may be some measurement error in self-reported age of first claiming, this comparison should is informative. The fraction of non-married males who claim at age 62 (38.6 percent) is only slightly higher than married (36.2 percent) and fraction of high school dropouts (37.0 percent) is lower than high school graduates (41.6 percent). In both cases, claiming differences do not seem to explain the large increases in mortality for males who were not married or did not complete high school. Note the claiming rates of white males are higher than non-white males, which may also explain why the mortality estimates for white males are larger and more precise than for non-white males.

Given that differences in Social Security claiming rates do not explain fully the mortality estimates, we next consider other changes associated with beginning to claim Social Security. In Appendix Table A5, we present estimates of age-62 changes in the fraction of respondents who report being partially and fully retired, working for pay and being without health insurance. HRS participants are interviewed approximately every two years, so we measure those outcomes in the last interview before turning 62 (i.e., at age 60 or 61) and in the first interview after turning 62 (i.e., at age 62 or 63). We also report the difference between the two. For males and females, we show these for age 62 claimants, for age 62 claimants with less than a high school education and for age 62 claimants who are not married. For comparison, we also present these for respondents not claiming Social Security at age 62.

There are large changes in each of these outcomes among age 62 claimants. For male age 62 claimants, the fraction partially or completely retired increases by 38.5 percentage points between surveys, the fraction working for pay decreases by 30.7 percentage points and the fraction without health insurance increases by 4.2 percentage points. These changes are substantially greater than for respondents not claiming Social Security at age 62. Interestingly, demographic differences in the changes in both retirement measures and working for pay are generally in line with the heterogeneity in mortality effects. For example, among males claiming at 62, the increase in partial or complete retirement is 37.6 percentage points for married and 42.0 percentage points for non-married males. Then, in terms of educational attainment, the change in retirement is 49.1 percentage points for male age 62 claimants who did not complete high school, 37.8 percentage points for high school graduates and 32.2 percentage points for college graduates. In contrast, female age 62 claimants report a 30.1 percentage point decrease in partial or complete retirement, while male and female non-claimants report changes of 8.6 and 11.2 percentage points, respectively. These results provide suggestive evidence that it may be major lifestyle changes that account for the increase in mortality at age 62, rather than Social Security per se.

6. Conclusion

Mortality is an important, well-measured, objective health outcome. The availability of administrative data on the entire U.S. population and a distinct Social Security eligibility age provides an opportunity to examine how Social Security (and related behaviors) affect health. We present evidence that there is an increase in mortality at age 62 for males of around two percent. This change is statistically significant and robust to different choices about how to implement a regression discontinuity design. The estimated increases are largest among males with low education or who were not married. While these demographic groups do not necessarily experience the largest rates of claiming Social Security at age 62, they do have the largest changes in terms of labor force exit and stopping work. The causes of death with the clearest increases at age 62 are accidents and two lung-related conditions: COPD and lung cancer.

The structure of the empirical approach means that we are estimating an effect that is local to age 62 and specific to age 62 claimants. It is not possible to establish whether age 62 Social Security claimants have elevated mortality over the longer term while they are receiving Social Security. It is also not possible to establish whether the mortality effects estimated here would apply to individuals who claim Social Security at later ages. Both of these are limitations and questions for future research.

7. References

Anderson, Michael L., Carlos Dobkin, and Tal Gross. 2014. The Effect of Health Insurance on Emergency Department Visits: Evidence from an Age-Based Eligibility Threshold. *Review of Economics and Statistics*, 96(1): 189-195.

Blau, David M., and Ryan Goodstein. 2010. Can Social Security Explain Trends in Labor Force Participation of Older Men in the United States? *Journal of Human Resources*. 45(2): 326-63.

Barecca, Alan I., Jason Lindo and Glen R. Waddell. 2015. Heaping-induced Bias in Regression-Discontinuity Designs. *Economic Inquiry*, forthcoming.

Behaghel, Luc, and David M. Blau. 2012. Framing Social Security Reform: Behavioral Responses to Changes in the Full Retirement Age. *American Economic Journal: Economic Policy*, 4(4): 1-30.

Behncke, Stefanie. 2012. Does Retirement Trigger Ill Health? *Health Economics.* 21(3): 282-300.

Blake, Hélène, and Clémentine Garrouste. 2013. Killing Me Softly: Work and Mortality among French Seniors. *University of York Health Econometrics and Data Group Working Paper 13/25*.

Bloemen, Hans, Stefan Hochguertel and Jochem Zweerink. 2013. The Causal Effect of Retirement and Mortality: Evidence from Targeted Incentives to Retire Early. *IZA Discussion Paper 7570.*

Bound, John, Michael Schoenbaum, Todd Stinebrickner, and Timothy Waidmann. 1999. The Dynamic Effects of Health on the Labor Force Transitions of Older Workers. *Labour Economics*, 6(2): 179-202.

Bound, John, and Timothy Waidmann. 2007. Estimating the Health Effects of Retirement. *Michigan Retirement Research Consortium Working Paper 2007-168*.

Calonico, Sebastian, Mattias Cattaneo, and Rocio Titiunik. 2014a. Robust Nonparametric Confidence Intervals for Regression-Discontinuity Designs. *Econometrica*, forthcoming.

Calonico, Sebastian, Mattias Cattaneo, and Rocio Titiunik. 2014b. Robust Data-Driven Inference in the Regression-Discontinuity Design. *Stata Journal*, forthcoming.

Card, David, Carlos Dobkin, and Nicole Maestas. 2009. Does Medicare Save Lives? *Quarterly Journal of Economics*, 124(2): 597–636.

Card, David, David S. Lee, Zhuan Pei and Andrea Weber. 2014. Local Polynomial Order in Regression Discontinuity Designs. *Brandeis University Economics Working Paper 2014/47*.

Charles, Kerwin K. 2004. Is Retirement Depressing?: Labor Force Inactivity and Psychological Well Being in Later Life. *Research in Labor Economics*, 23: 269-299.

Coe, Norma B., and Maarten Lindeboom. 2008. Does Retirement Kill You? The Evidence from Early Retirement Windows. *IZA Discussion Paper No. 3817*.

Coe, Norma B., and Gema. Zamarro. 2011. Retirement Effects on Health in Europe. *Journal of Health Economics*, 30(1): 77-86.

Dave, Dhaval, Inas Rashad, and Jasmina Spasojevic. 2008. The Effects of Retirement on Physical and Mental Health Outcomes. *Southern Economic Journal*, 75(2): 497-523.

Engelhardt, Gary, and Anil Kumar. 2014. Taxes and the Labor Supply of Older Americans: Recent Evidence from the Social Security Earnings Test. *National Tax Journal*, 67(2): 443-458.

Evans, William N., and Timothy J. Moore. 2011. The Short-term Mortality Consequences of Income Receipt. *Journal of Public Economics*, 95(11-12): 1410-1424.

French, Eric. 2005. The Effects of Health, Wealth, and Wages on Labor Supply and Retirement Behavior. *Review of Economic Studies*, 72(2): 395-427.

Friedberg, Leora. 2000. The Labor Supply Effects of the Social Security Earnings Test. *Review of Economics and Statistics*, 82(1): 48-63.

Gelber, Alexander, Damon Jones and Daniel Sacks. 2013. Earnings Adjustment Frictions: Evidence from the Social Security Earnings Test. *NBER Working Paper 19491*.

Gelman, Andrew and Guido Imbens, 2014. Why High-order Polynomials Should Not Be Used in Regression Discontinuity Designs. *NBER Working Paper 20405.*

Gustman, Alan, and Thomas Steinmeier. 2005. The Social Security Early Retirement Age in a Structural Model of Retirement and Wealth. *Journal of Public Economics*, 89(2-3): 441-463.

Hernaes, Erik, Simen Markussen, John Pigott and Ola Vestad 2013. Does Retirement Age Impact Mortality? *Journal of Health Economics*, 32: 586-598.

Hu, Yuanreng, and Noreen Goldman. 1990. Mortality Differentials by Marital Status: An International Comparison. *Demography*, 27(2): 233-250.

Insler, Michael. 2014. The Health Consequences of Retirement. *Journal of Human Resources*, 49(1): 195-233.

Kane, Thomas J. 2003. A Quasi-experimental Estimate of the Impact of Financial Aid on College-Going. *NBER Working Paper 9703*.

Kuhn, Andreas, Jean-Philippe Wuellrich and Josef Zweimuller. 2010. Fatal Attraction? Access to Early Retirement and Mortality. *IZA Discussion Paper 5160*.

Lantz, Paula, James S. House, James M. Lepkowski, David R. Williams, Richard P. Mero and Jieming Chen. 1998. Socioeconomic Factors, Health Behaviors, and Mortality: Results from a Nationally Representative Prospective Study of US Adults. *Journal of American Medical Association*, 279(21): 1703-1708.

Lee, David S. and Thomas Lemieux. 2010. Regression Discontinuity Designs in Economics. *Journal of Economic Literature* 48: 281-355.

Li, Xiaoyan, Michael Hurd, and David S. Loughran. 2008. The Characteristics of Social Security Beneficiaries Who Claim Benefits at the Early Entitlement Age. *AARP Public Policy Institute Research Report No. 2008-19*. Washington, DC: AARP.

Mastrobuoni, Giovanni. 2009. Labor Supply Effects of the Recent Social Security Benefit Cuts: Empirical Estimates Using Cohort Discontinuities. *Journal of Public Economics*, 93(11-12): 1224-33.

McGarry, Kathleen. 2004. Health and Retirement: Do Changes in Health Affect Retirement Expectations? *Journal of Human Resources*, 39(3): 624-648.

Neuman, Kevin. 2008. Quit Your Job and Get Healthier? The Effect of Retirement on Health. *Journal of Labor Research*, 29(2): 177-201.

Rust, John, and Christopher Phelan. 1997. How Social Security and Medicare affect Retirement Behavior in a World of Incomplete Markets. *Econometrica*, 65(4): 781-831.

Snyder, Stephen E., and William N. Evans. 2006. The Effect of Income on Mortality: Evidence from the Social Security Notch. *Review of Economics and Statistics*, 88(3): 482–495.

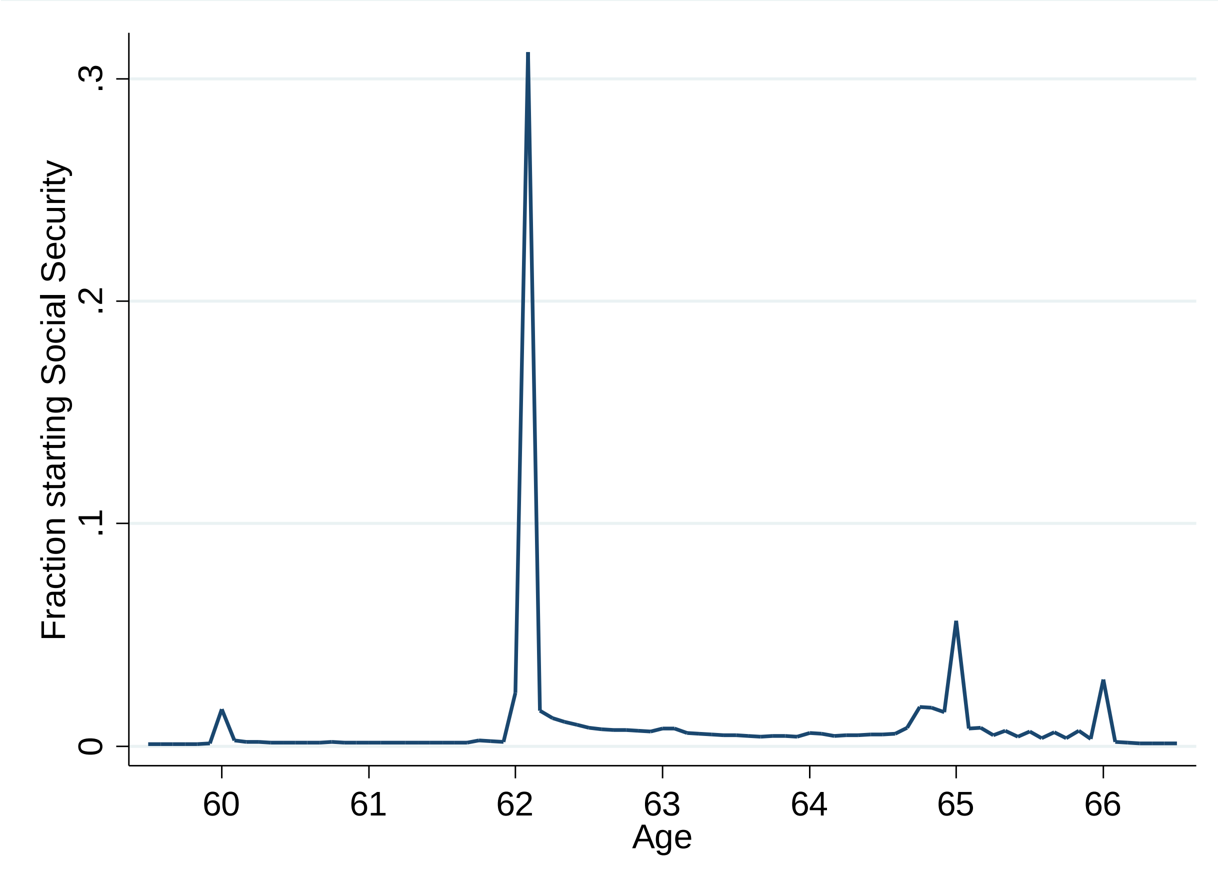
Social Security Administration. 2012. *Annual Statistical Supplement to the Social Security Bulletin, 2011*. Washington DC: SSA Office of Retirement and Disability Policy.

Song, Jae, and Joyce Manchester. 2008. Have People Delayed Claiming Social Security Benefits? Responses to Changes in Social Security Rules. *Social Security Bulletin*, 67(2): 1-23.

Sorlie, Paul D., and Norman J. Johnson. 1996. Validity of Education Information on the Death Certificate. *Epidemiology*, 7: 437–439.

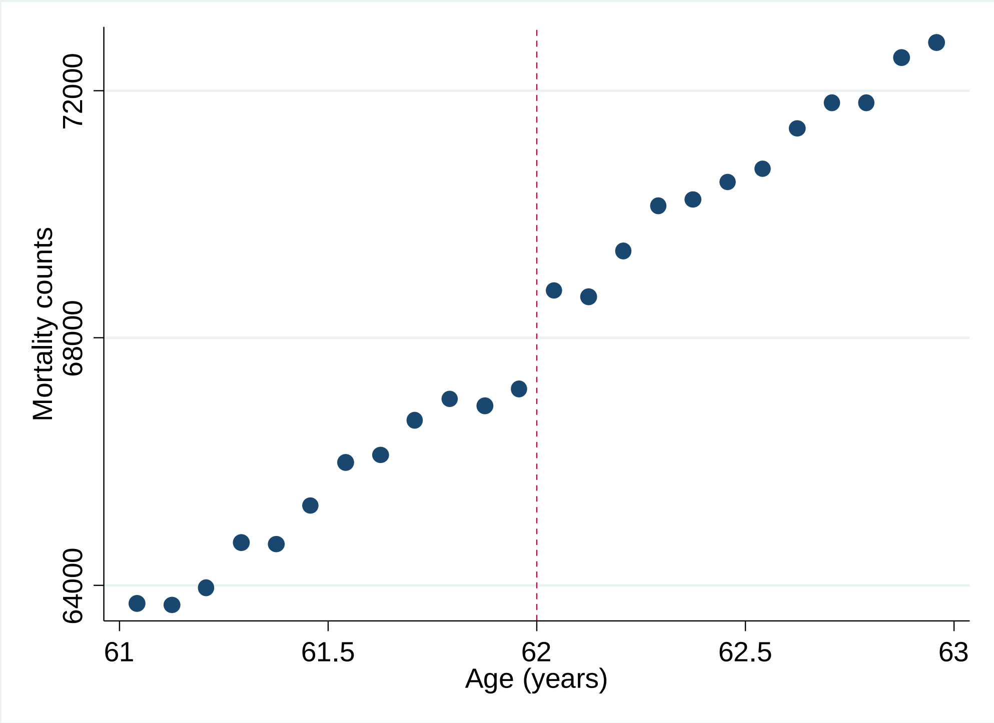
Sorlie, Paul D., Eugene Rogot, and Norman J. Johnson. 1992. Validity of Demographic Characteristics on the Death Certificate. *Epidemiology*, 3(2): 181-184.

Figure 1 Rate of New Social Security Claims, Ages 59 to 67



Notes: We use birth cohorts from 1921 to 1948 and include new claims by both workers and dependents for the Disability, Retirement and Survivors components of Social Security. This is converted to rates based on population estimates from the Current Population Survey.

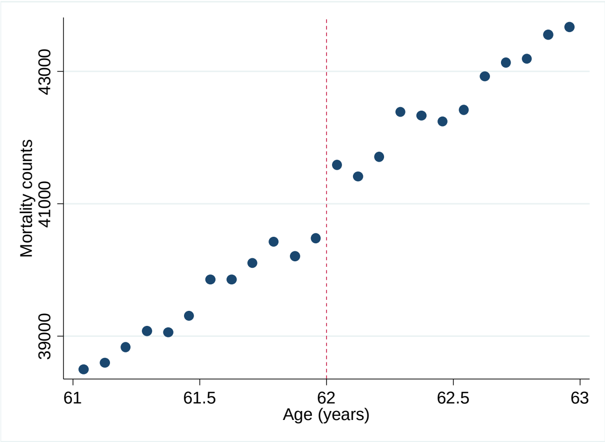
Figure 2 Monthly Mortality Counts in Relation to Turning Age 62, Cohorts Born 1921-1948



Notes: Data is from restricted-use versions of the Multiple Cause of Death data held at the National Center for Health Statistics. Reported mortality counts are for the 1921 to 1948 cohorts. The figures show the number of deaths by age measured in months.

Figure 3 Monthly Mortality Counts in Relation to Turning Age 62, By Sex

A: Males



B: Females

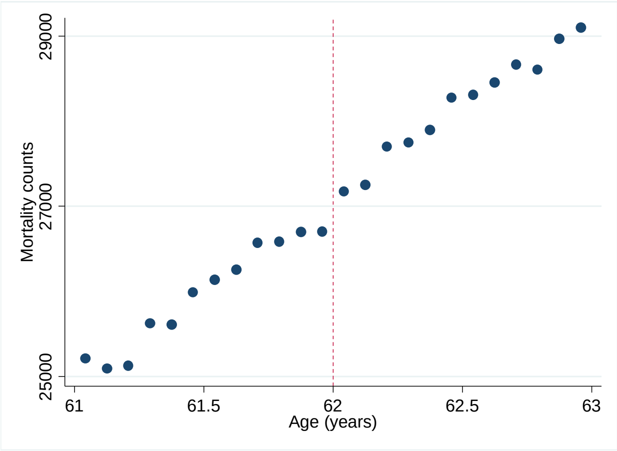


Table 1 Summary Statistics, Multiple Cause of Death Data

	All		Males		Females	
	Age 61	Age 62	Age 61	Age 62	Age 61	Age62
Total deaths	785,871	848,728	474,272	510,561	311,599	338,167
Ave. deaths per month	65,489	70,727	39,523	42,547	25,967	28,181
Race						
White (%)	82.7	82.8	83.4	83.5	81.6	81.7
Black (%)	15.3	15.1	14.7	14.5	16.2	16.1
Other race (%)	2.0	2.1	1.9	2.0	2.2	2.2
Marital Status						
Single (%)	9.0	8.9	10.0	9.8	7.5	7.5
Married (%)	60.2	59.5	65.2	65.0	52.5	51.3
Divorced (%)	18.7	18.6	18.9	18.8	18.5	18.4
Widowed (%)	12.1	13.0	5.9	6.4	21.4	22.9
Educational Attainment						
Less than high school (%)	26.4	26.8	27.4	27.6	25.1	25.6
High school graduate (%)	59.3	59.0	56.7	56.5	63.1	62.8
College graduate (%)	14.3	14.2	16.0	15.9	11.8	11.7
Place of Death						
Hospital (%)	59.3	58.6	59.3	58.6	59.4	58.6
Nursing home/institution (%)	6.8	7.4	6.0	6.5	8.0	8.7
Residence or other location (%)	33.9	34.1	34.8	35.0	32.6	32.7
Cause of Death						
Heart attacks (%)	17.3	17.5	19.4	19.5	14.1	14.5
Other heart disease (%)	10.5	10.1	12.2	11.6	7.9	7.8
Stroke (%)	4.0	4.1	3.6	3.7	4.6	4.6
COPD (%)	4.4	4.8	4.0	4.3	5.1	5.4
Lung cancer (%)	12.5	12.6	13.0	13.1	11.6	11.7
Breast cancer (%)	3.1	3.0	0.0	0.0	7.9	7.4
Leukemia (%)	1.0	1.0	1.0	1.0	1.0	1.0
Other cancers (%)	20.1	20.0	19.5	19.6	20.9	20.6
External causes (%)	4.5	4.1	5.3	4.8	3.3	3.1
All other causes (%)	22.7	22.9	22.0	22.2	23.6	24.0

Notes: Data is from restricted-use versions of the Multiple Cause of Death data held at the National Center for Health Statistics. The sample includes all deaths for cohorts born between 1921 and 1948. Marital status is missing in 0.9 percent of cases and place of death is missing in 0.7 percent of cases. As educational attainment is only available after 1989, it is available for 69.1 percent of all cases. The category "Residence or other location" includes those classified as "Dead on arrival" at the hospital.

Table 2 Regression Estimates of Increase in Mortality at Age 62

Regression type	All (1)	Males (2)	Females (3)
Global parametric regressions (bandwidth = 12 months)			
Quadratic regression	0.0135***	0.0185***	0.0058
	(0.0043)	(0.0049)	(0.0049)
Cubic regression	0.0197***	0.0236***	0.0138***
	(0.0049)	(0.0060)	(0.0047)
Quartic regression	0.0193***	0.0243***	0.0116***
	(0.0051)	(0.0082)	(0.0043)
Polynomial minimizing AICc	Cubic	Quadratic	Quartic
Local nonparametric regressions			
Local linear using data-driven bandwidth	0.0142***	0.0215***	0.0103***
	(0.0036)	(0.0041)	(0.0030)
Data-driven bandwidth	10 months	7 months	6 months
Local quadratic using data-driven bandwidth	0.0194***	0.0233***	0.0131***
	(0.0039)	(0.0058)	(0.0026)
Data-driven bandwidth	7 months	7 months	8 months

Notes: ** denotes $p<0.05$, *** denotes $p<0.01$. Data is from restricted-use versions of the Multiple Cause of Death data held at the National Center for Health Statistics. The global parametric regressions allow for the polynomial to vary either side of the discontinuity and we report robust standard errors. AICc is the corrected (finite-sample) Akaike Information Criteria. The nonparametric regressions are estimated using the "robust data-driven" procedures of Calonico, Cattaneo and Titiunik (2014). We use a triangular kernel, robust standard errors, and their bandwidth selection and bias correction procedures. See text for more details.

Table 3 Estimated Change in Mortality at Age 62, Male Subgroups

	Local linear (1)	Global quadratic (2)	Fraction deaths (3)
Marital status			
Married	0.0130**	0.0081	65.1%
	(0.0058)	(0.0057)	
	6 months	12 months	
Not married	0.0415***	0.0377***	34.9%
	(0.0079)	(0.0107)	
	6 months	12 months	
- Single	0.0558***	0.0514***	9.9%
	(0.0056)	(0.0111)	
	6 months	12 months	
- Divorced	0.0305***	0.0337**	18.9%
	(0.0111)	(0.0137)	
	8 months	12 months	
- Widowed	0.0330	0.0262	6.2%
	(0.0202)	(0.0248)	
	11 months	12 months	
Educational attainment			
Did not complete high school	0.0303***	0.0275**	27.6%
	(0.0103)	(0.0115)	
	8 months	12 months	
Completed high school, not college	0.0087	0.0099	56.5%
	(0.0050)	(0.0066)	
	7 months	12 months	
Completed college	0.0146	0.0187	15.9%
	(0.0146)	(0.0181)	
	8 months	12 months	
Race			
White	0.0237***	0.0209***	83.5%
	(0.0042)	(0.0053)	
	7 months	12 months	
Non-white	0.0078	0.0068	16.5%
	(0.0104)	(0.0115)	
	8 months	12 months	

	Local linear (4)	Global quadratic (5)	Fraction deaths (6)
Place of death			
Out of hospital/ institution	0.0339***	0.0323***	34.6%
	(0.0074)	(0.0073)	
	12 months	12 months	
In hospital	0.0124*	0.0130*	58.5%
	(0.0064)	(0.0070)	
	11 months	12 months	
In nursing home/ institution	-0.0022	-0.0079	6.2%
	(0.0161)	(0.0194)	
	8 months	12 months	
Cause of death			
Heart and lung conditions	0.0250**	0.0135	39.2%
	(0.0106)	(0.0119)	
	6 months	12 months	
- Heart attacks	0.0159	0.0072	19.4%
	(0.0138)	(0.0172)	
	8 months	12 months	
- COPD	0.0696***	0.0496**	4.2%
	(0.0089)	(0.0180)	
	7 months	12 months	
- Not heart attacks or COPD	0.0064	0.0118	15.6%
	(0.0106)	(0.0115)	
	7 months	12 months	
Cancers	0.0262***	0.0263***	33.6%
	(0.0072)	(0.0084)	
	8 months	12 months	
- Lung cancer	0.0531***	0.0510***	13.1%
	(0.0097)	(0.0108)	
	7 months	12 months	
- Not lung cancer	0.0099	0.0106	20.6%
	(0.0080)	(0.0092)	
	7 months	12 months	
External causes	0.0314***	0.0399**	5.0%
	(0.0103)	(0.0163)	
	7 months	12 months	
All other causes	0.0113	0.0109	22.1%
	(0.0109)	(0.0106)	
	6 months	12 months	

Notes: ** denotes p<0.05, *** denotes p<0.01. For each subgroup, we show the coefficient, standard error and bandwidth used for both the local linear regression (with CCT bandwidth) and global quadratic regression (with 12-month bandwidth). Fraction of deaths is based on deaths at ages 61 and 62. See the notes to Table 2 for more details.

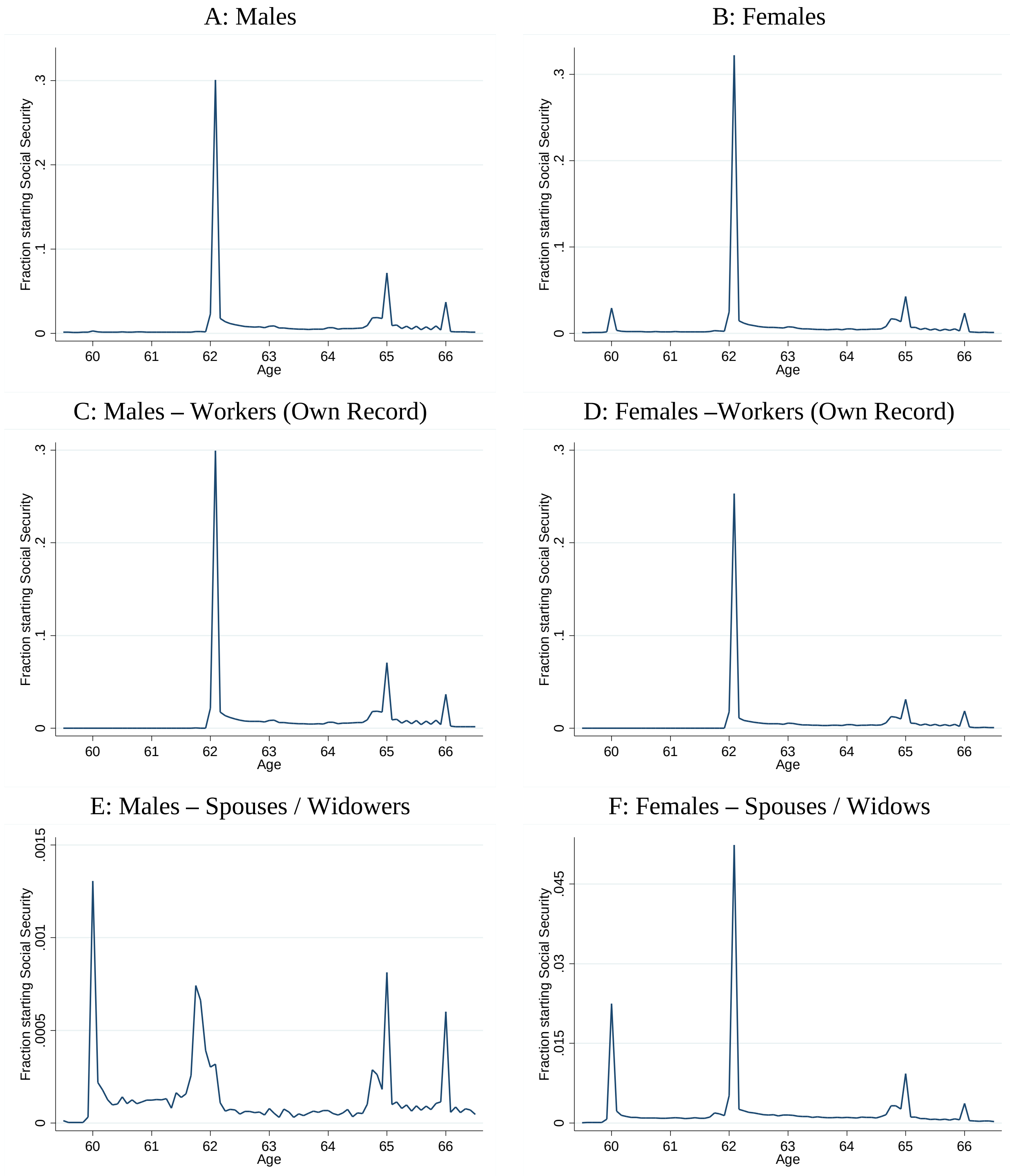
Table 4 Estimates for Different Cohorts of Males

	Local linear (1)	Global quadratic (2)
Cohorts with a FRA equal to 65 years: 1921-1937	0.0148	0.0170**
	(0.0090)	(0.0068)
	11 months	12 months
Cohorts with a FRA greater than 65 years: 1938-1948	0.0244**	0.0179***
	(0.0099)	(0.0069)
	8 months	12 months
- Cohorts with a FRA between 65 & 66 years: 1938-1942	0.0259	0.0190
	(0.0178)	(0.0151)
	10 months	12 months
- Cohorts with a FRA equal to 66 years: 1943-1948	0.0232**	0.0170***
	(0.0098)	(0.0062)
	7 months	12 months

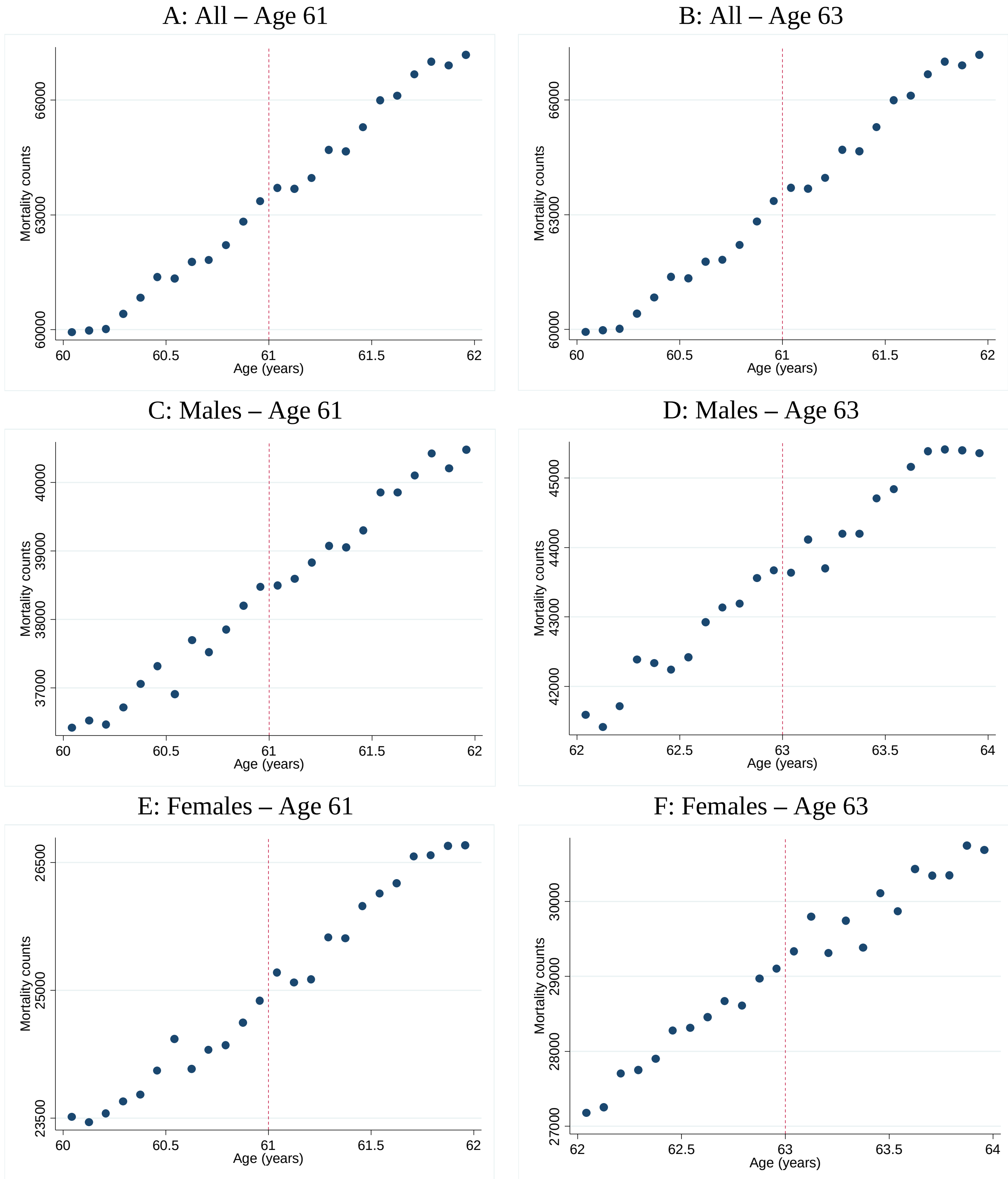
Notes: ** denotes p<0.05, *** denotes p<0.01. FRA is “Full Retirement Age.” For each subgroup, we show the coefficient, standard error and bandwidth used for both the local linear regression (with CCT bandwidth) and global quadratic regression (with 12-month bandwidth). See the notes to Table 2 for more details.

ONLINE APPENDICES

Appendix Figure A1 Social Security Claiming Rates by Sex and Types of Claim, Ages 59 to 67

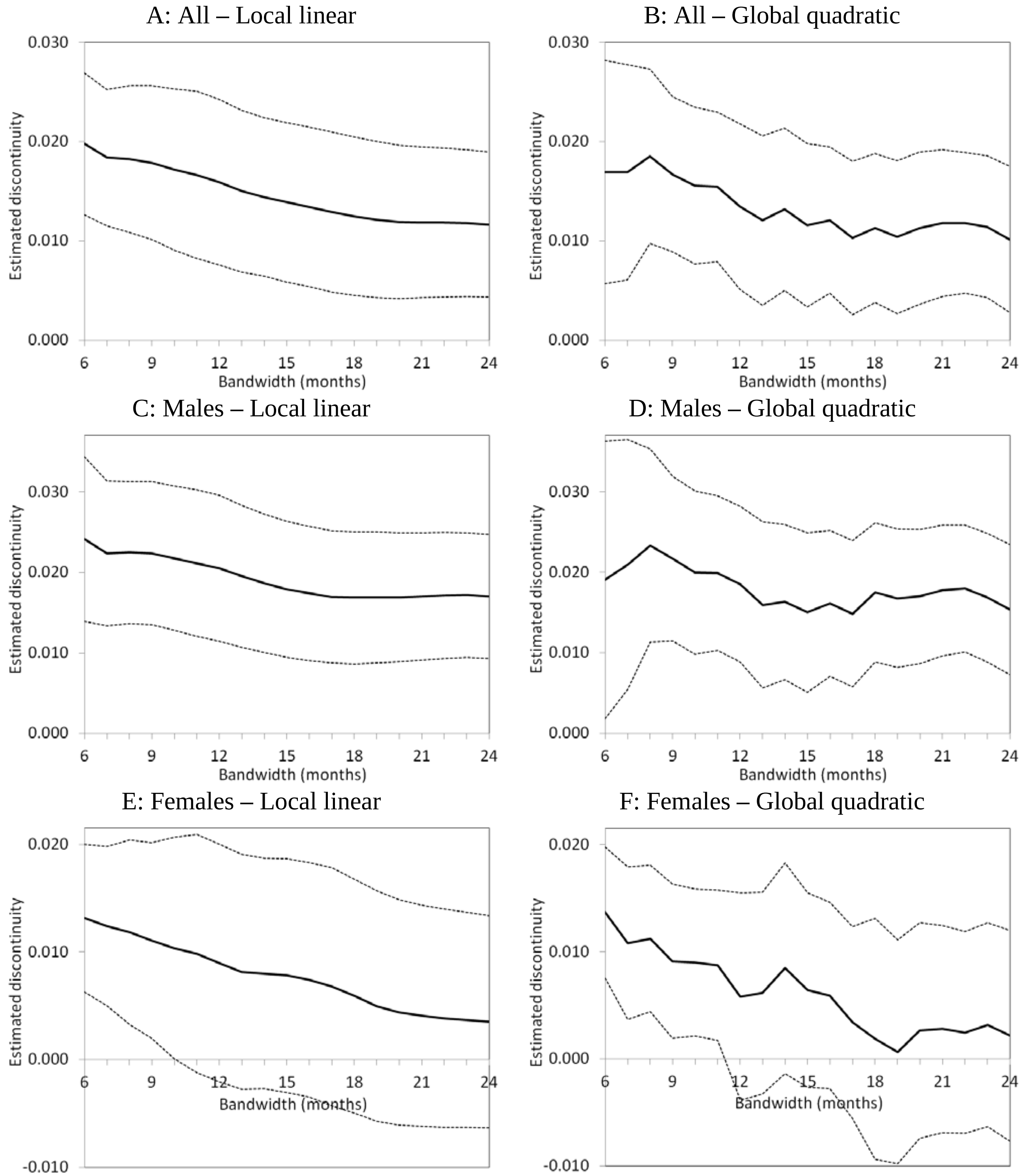


Appendix Figure A2 Monthly Mortality Counts in Relation to Turning Age 61 and 63



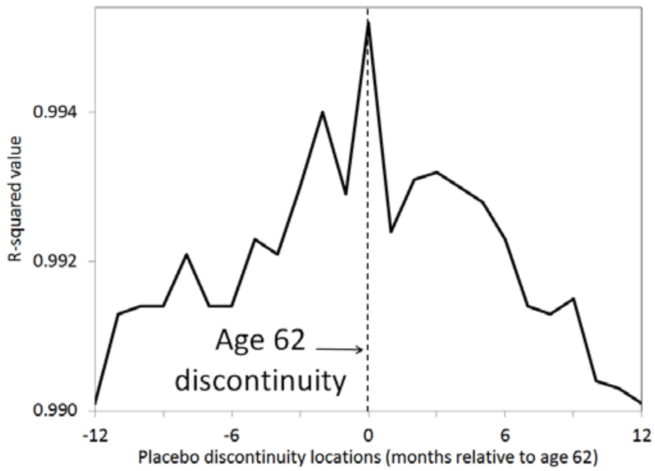
Notes: Data is from restricted-use versions of the Multiple Cause of Death data held at the National Center for Health Statistics. Reported mortality counts are for the 1921 to 1948 cohorts. The figures report the number of deaths by age measured in months.

Appendix Figure A3 Robustness of Estimates to Bandwidth

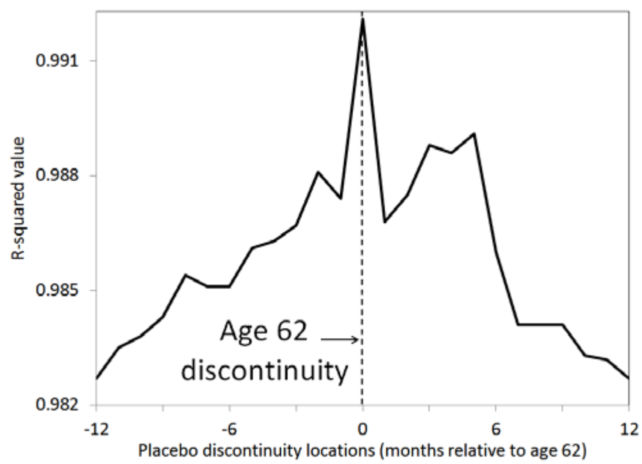


Appendix Figure A4 R-squared as Function of Location of the Discontinuity, Global Quadratic

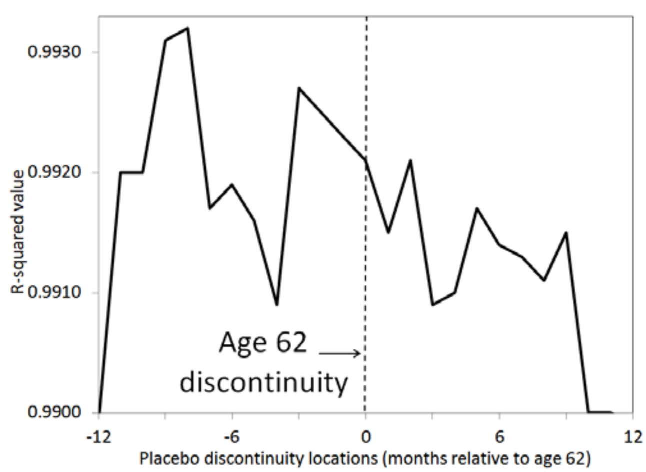
A: All



B: Males

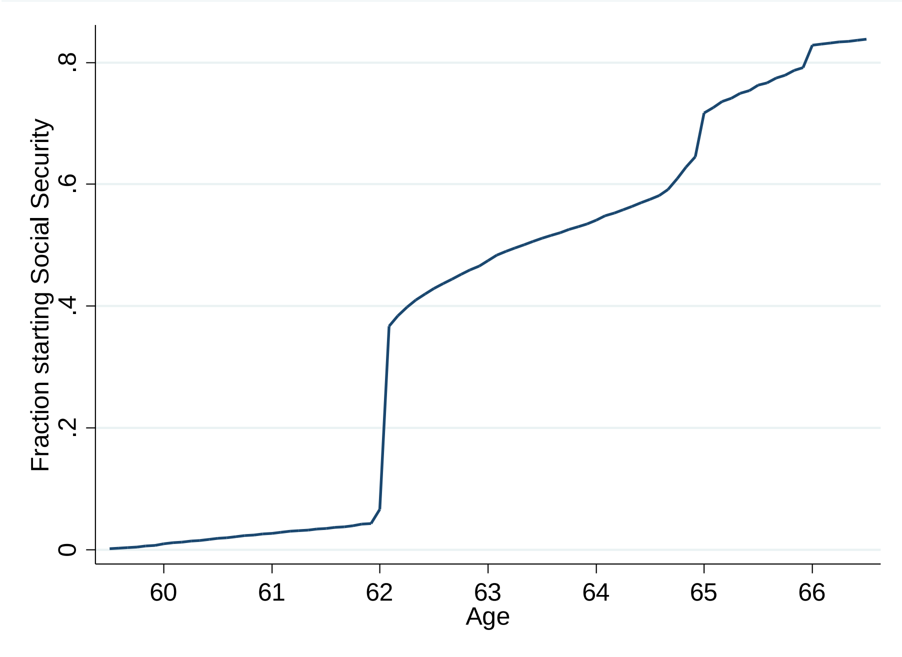


C: Females

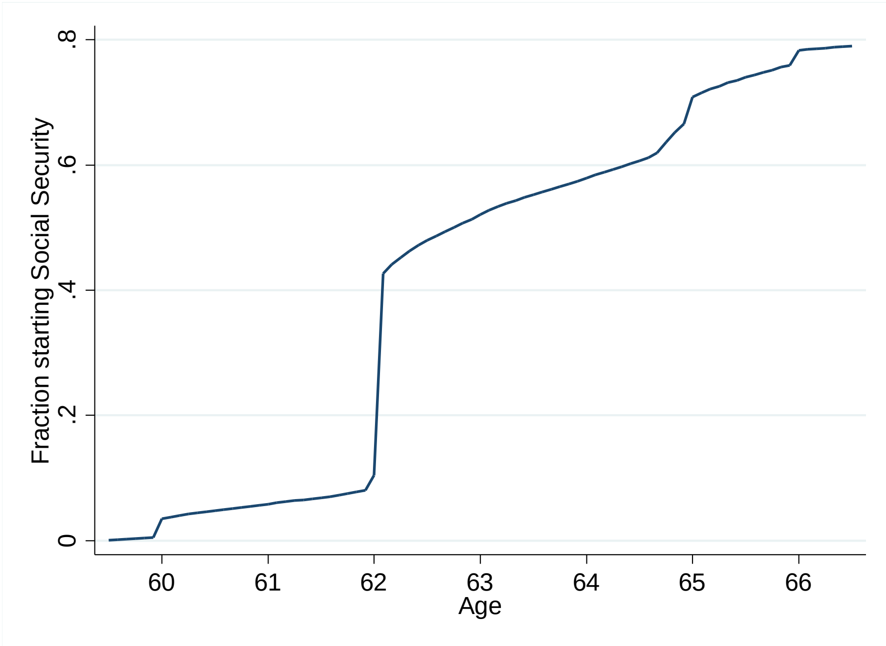


Appendix Figure A5 Cumulative Social Security Claiming Rates

A: Males



B: Females



Appendix Table A1 Robustness of Regression Estimates

Regression type		Extra controls		Unit of observation		Using SS
	Main estimate	Year of birth FE	Month of death FE	Daily data	Weekly data	eligibility date
	(1)	(2)	(3)	(4)	(5)	(6)
Global parametric regressions (bandwidth = 12 months)						
Global quadratic	0.0135***	0.0132***	0.0133***	0.0132***	0.0137**	0.0134
	(0.0043)	(0.0039)	(0.0041)	(0.0049)	(0.0067)	(0.0084)
	12 months	12 months	12 months	365 days	52 weeks	12 months
Global cubic	0.0197***	0.0195***	0.0213***	0.0204***	0.0200**	0.0254***
	(0.0049)	(0.0043)	(0.0039)	(0.0067)	(0.099)	(0.0068)
	12 months	12 months	12 months	365 days	52 weeks	12 months
Global quartic	0.0193***	0.0187***	0.0208***	0.0224***	0.0235**	0.0333***
	(0.0051)	(0.0043)	(0.0041)	(0.0081)	(0.0114)	(0.0065)
	12 months	12 months	12 months	365 days	52 weeks	12 months
Local nonparametric regressions						
Local linear	0.0142***	--	--	0.0135***	0.0223**	0.0166**
	(0.0036)			(0.0047)	(0.0105)	(0.0074)
CCT bandwidth	10 months			305 days	18 weeks	8 months
Local quadratic	0.0194***	--	--	0.0221***	0.0239**	0.0359***
	(0.0039)			(0.0077)	(0.0115)	(0.0043)
CCT bandwidth	7 months			218 days	27 weeks	6 months

Notes: * denotes p<0.10, ** denotes p<0.05, *** denotes p<0.01. Data is from restricted-use versions of the Multiple Cause of Death data held at the National Center for Health Statistics. The global parametric regressions allow for the polynomial to vary either side of the discontinuity and we report robust standard errors. AICc is the corrected (finite-sample) Akaike Information Criteria. The nonparametric regressions are estimated using the “robust data-driven” procedures of Calonico, Cattaneo and Titiunik (2014). We use a triangular kernel, robust standard errors, and their bandwidth selection and bias correction procedures. See text for more details.

Appendix Table A2 Robustness of Regression Estimates and Placebo Tests at Ages 61 and 63

	All		Males		Females	
Discontinuity at:	Local linear (7)	Global quadratic (7)	Local linear (8)	Global quadratic (8)	Local linear (8)	Global quadratic (8)
Age 61	-0.0023	-0.0044	-0.0058**	-0.0076**	0.0105	0.0084
	(0.0029)	(0.0047)	(0.0025)	(0.0030)	(0.0059)	(0.0069)
	5 months	12 months	7 months	12 months	7 months	12 months
Age 63	-0.0023	-0.0044	-0.0061	-0.0088	0.0060	0.0006
	(0.0029)	(0.0047)	(0.0035)	(0.0049)	(0.0073)	(0.0085)
	5 months	12 months	9 months	12 months	7 months	12 months

Notes: ** denotes p<0.05, *** denotes p<0.01. Data is from restricted-use versions of the Multiple Cause of Death data held at the National Center for Health Statistics. The global parametric regressions allow for the polynomial to vary either side of the discontinuity and we report robust standard errors. AICc is the corrected (finite-sample) Akaike Information Criteria. The nonparametric regressions are estimated using the "robust data-driven" procedures of Calonico, Cattaneo and Titiunik (2014). We use a triangular kernel, robust standard errors, and their bandwidth selection and bias correction procedures. See text for more details.

Appendix Table A3 Estimated Increases in Social Security Claiming at Age 62

	Males (1)	Females (2)
Global parametric (bandwidth=12 months		
Quadratic regression	0.3111***	0.3351***
	(0.0032)	(0.0024)
Cubic regression	0.3025***	0.3282***
	(0.0010)	(0.0009)
Quartic regression	0.3012***	0.3275***
	(0.0008)	(0.0008)
Local nonparametric		
Linear regression	0.3068***	0.3324***
	(0.0033)	(0.0027)
B/width	6 months	7 months
Quadratic regression	0.3015***	0.3276***
	(0.0014)	(0.0011)
B/width	7 months	7 months

Notes: ** denotes p<0.05, *** denotes p<0.01. Social Security claiming data are from tabulations of the Master Beneficiary Record (1 percent extract). The global parametric regressions allow for the polynomial to vary either side of the discontinuity and we report robust standard errors. The nonparametric regressions are estimated using the "robust data-driven" procedures of Calonico, Cattaneo and Titiunik (2014). We use a triangular kernel, robust standard errors, and their bandwidth selection and bias correction procedures. See text for more details.

Appendix Table A4 Males' Social Security Starting Ages by Characteristics at Age 60 or 61

	First claimed Social Security at:					Fraction of sample
	<=61	62	63- <FRA	≥FRA	Other	
Marital status						
Married (%)	14.7	36.2	20.0	14.3	14.8	82.4
Non-married (%)	21.0	38.6	15.6	10.7	14.1	17.6
- Single (%)	19.8	35.1	17.4	10.3	17.3	6.3
- Divorced (%)	20.4	39.6	14.6	11.5	13.9	8.7
- Widowed (%)	26.8	43.1	15.3	8.5	6.4	2.6
Educational attainment						
< High school (%)	30.3	37.0	13.4	9.3	9.9	22.6
Completed high school (%)	14.8	41.6	19.1	11.6	12.9	52.6
Completed college (%)	9.1	27.5	22.8	19.8	20.9	24.8
Race						
White (%)	14.8	37.1	20.0	14.1	14.1	83.0
Non-white (%)	24.5	33.6	13.0	10.0	18.9	17.0
Sample size	654	1,502	784	557	600	4,097

Notes: These data are from the Health and Retirement Study (v.N of RAND compilation), which incorporates data from eleven waves and covers1992 to 2012. The characteristics are from the last survey prior to turning age 62. The tabulations are based on person-specific weights. FRA stands for the "Full Retirement Age."

Appendix Table A5 Changes in Retirement, Paid Work and Health Insurance at Age 62

	Last interview before 62 (1)	First interview after 62 (2)	Difference [(2)-(1)] (3)
	Percent Who Say are Partially or Fully Retired		
SS at 62: males (%)	45.5	84.0	38.5
SS at 62: females (%)	39.1	69.2	30.1
SS at 62: males, married (%)	45.6	83.2	37.6
SS at 62: males, non-married (%)	45.2	87.2	42.0
SS at 62: males, <high school (%)	30.3	79.4	49.1
SS at 62: males, high school grad. (%)	46.9	84.7	37.8
SS at 62: males, college graduate (%)	53.7	85.9	32.2
Other respondents: males (%)	33.6	40.5	6.9
Other respondents: females (%)	32.7	42.7	10.0
	Percent Who are Working for Pay		
SS at 62: males (%)	65.6	34.9	-30.7
SS at 62: females (%)	49.2	28.3	-20.9
SS at 62: males, married (%)	66.8	36.7	-30.1
SS at 62: males, non-married (%)	61.0	28.3	-32.7
SS at 62: males, <high school (%)	71.8	34.6	-37.2
SS at 62: males, high school grad. (%)	65.5	35.3	-30.2
SS at 62: males, college graduate (%)	60.8	34.2	-26.6
Other respondents: males (%)	67.8	63.0	-4.8
Other respondents: females (%)	55.5	50.0	-5.5
	Percent without Health Insurance		
SS at 62: males (%)	17.4	21.6	4.2
SS at 62: females (%)	22.2	27.1	4.9
SS at 62: males, married (%)	14.0	19.7	5.7
SS at 62: males, non-married (%)	30.3	28.7	-1.6
SS at 62: males, <high school (%)	29.1	35.8	6.7
SS at 62: males, high school grad. (%)	14.0	17.5	3.5
SS at 62: males, college graduate (%)	17.0	21.4	4.4
Other respondents: males (%)	13.2	11.7	-1.5
Other respondents: females (%)	17.1	17.7	0.6

Notes: Data are from the Health and Retirement Study (v.N of RAND compilation), which incorporates data from eleven waves and covers1992 to 2012. The tabulations are based on person-specific weights.