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Abstract

Reductions in the implicit taxation of Social Security benefits from reducing or eliminating the Retirement Earnings Test (RET) are an appealing – and in many cases successful – means of encouraging labor supply of older individuals receiving benefits. The downside, however, is that the same policy reforms can encourage earlier claiming of Social Security benefits, which permanently lowers benefits paid in the future. Depending on the magnitude of the effects on earnings and how households or individuals adjust their consumption and savings decisions, the net effect can be lower incomes at much older ages well beyond when people have retired. We explore the consequences of the 2000 reforms eliminating the RET from the Full Retirement Age to age 69 for the longer-run evolution of income, focusing in particular on the incidence of low income among older women, who are more likely to have become dependent mainly on income from their Social Security benefits. We find that the elimination of the RET increased the likelihood of having low incomes among women in their mid-70s and older – ages at which the lower benefits from claiming earlier outweigh possibly higher income in the period when women or their husbands increased their labor supply.

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I. Introduction

The Social Security retirement earnings test (RET) establishes a threshold for which wage earnings in excess of the threshold cause a reduction in benefits. Prior to 2000, the RET reduced benefits by $1 for every $3 earned above the threshold, for beneficiaries from the Full Retirement Age (FRA) to age 69 (Social Security Administration, 2010). Beneficiaries age 62 to the FRA are subject to a more restrictive RET, reducing benefits by $1 for every $2 earned above the threshold, and a lower threshold. The Senior Citizens' Freedom to Work Act of 2000 repealed the RET for beneficiaries who had attained the FRA, as well as making the RET less stringent in the year an individual reached the FRA. There were earlier reforms of the RET in the 1980s.

The RET is viewed as a tax (Liebman and Luttmer, 2011) despite the fact that the lost benefits are provided in the future. As a result, the predicted effect of the removal of the RET on labor supply (and hence earnings) depends on where the beneficiary was on the budget constraint prior to the elimination of the RET (Haider and Loughran, 2008). Figure 1 provides the budget constraints with and without the RET. Beneficiaries choosing low hours prior to 2000 (located on segment DC in Figure 1) should be unaffected by the elimination of the RET. Among those affected, beneficiaries whose labor supply led to the elimination of all their Social Security benefits (located on segment BE in Figure 1) experience an income effect, hence likely decreasing their hours worked. Those bunching just at or below the RET threshold (point D in Figure 1) should increase their hours worked through the substitution effect. And for those receiving reduced benefits due to earnings above the threshold (located on segment DE), there is both an income and a substitution effect, so the net effect is ambiguous.

The average effect of the 2000 removal on beneficiaries' labor supply decision and earnings is unclear from the theoretical predictions. But earlier research on either primary beneficiaries or only men finds bunching of earnings just below levels where the RET applies (Friedberg, 2000; Gruber and Orszag, 2003; Song and Manchester, 2007; Haider and Loughran, 2008), and increases in labor supply and earnings from eliminating the threshold. Recent work focused on the 2000 reforms (Figinski, 2013) studies female primary and spousal beneficiaries, and finds that, like men, female primary beneficiaries affected by the removal of the RET increased their labor supply and earnings, although there was no change for spousal beneficiaries.[1]

Concomitant with changes in labor supply, however, are effects on benefit claiming and hence on benefit levels. Men or women working enough to be subject to the RET have an incentive to delay claiming, and conversely to claim earlier when the RET is eliminated or reduced, resulting in lower monthly Social Security benefits. In particular, with the 2000 policy changes, there should be some people who move the age of claiming benefits from above the FRA down to the FRA, because the RET was eliminated for those attaining the FRA. Looking at reforms in the 1980s, Gruber and Orszag (2003) find this for men, and Figinski (2013) also finds that women – both primary and spousal beneficiaries – claimed benefits earlier when the RET was eliminated in 2000, for certain age groups. (See also Song and Manchester, 2007.) Those who claim benefits early and continue to work will be able to supplement their wage income with Social Security benefits. Yet these individuals may not save for the time when they can no longer work and thus may have lower incomes – including Social Security payments – than if they had not claimed benefits earlier in response to the elimination of the RET. As explained below, this problem may be particularly severe for older women.

[1] The absence of a labor supply response among female spousal beneficiaries could reflect low skills, or careers where it is difficult to re-enter the labor market or increase earnings, as well as an income effect in response to husband's higher earnings stemming from removal of the RET

As a consequence of these responses, removing the RET could have the unintended effect of increasing poverty – or the incidence of low-income more generally – among much older individuals. Gruber and Orszag (2003) were the first to point this out, suggesting that the removal of the RET for younger beneficiaries "could produce a nontrivial increase in elderly poverty" (p. 771). Of course there is a potential offsetting response if the elimination of the RET implies that people work longer and, therefore, run down their assets more slowly with positive implications for resources available at older ages.

It is this potential for longer-term response of changes in the RET that we explore in this research. We know from past work (e.g., Sandell and Iams, 1997) that old-age poverty is common among women beginning around age 75, and that this may stem in part from how Social Security benefits change with early claiming, so that women, who generally outlive men, will face actuarially unfair benefit reductions from early claiming (by them or their husbands) – a problem that is exacerbated by the structure of spousal and survivor benefits (Sass et al., 2007).[2]

We focus on women, in particular, because they generally outlive men. Hence, women – and especially older women – are more likely to reach the point where reliance on Social Security benefits is higher. In this case, the lower benefits from claiming earlier are more likely to outweigh the higher earnings or saving from increased labor supply in response to the elimination of the RET.

Although the research, of necessity, looks backward at responses to past changes in the RET, it has potential implications for the future if policymakers were to eliminate or further reduce the RET at age 62 in an effort to encourage more work beyond age 62. The positive labor supply response of female primary beneficiaries, coupled with the results from the previous literature showing labor supply increases of men and primary beneficiaries (regardless of sex),

[2] For example, husbands who are much older than their wives may not take account of the increase in the value of survivors' benefits for their wives from delayed claiming.

may suggest that, as the FRA continues to rise, the RET should be removed for all beneficiaries.[3] Yet as the FRA increases further (beginning with the 1955 birth cohort), individuals who claim benefits at age 62 will receive a greater reduction in benefits for claiming benefits early. For those who claim benefits early and do not continue to work, the lower benefit level will make it more difficult to make ends meet. Those who claim benefits early and continue to work will be able to supplement their wage income with Social Security benefits. But these individuals may not save for the time when they can no longer work and become dependent on Social Security payments for their income. Thus, creating incentives to work more by eliminating the RET – because it also increases incentives to claim benefits at an earlier age – may increase poverty or the incidence of low-income more generally among older individuals, and especially older women.

II. Past Research

The earnings increases that Figinski finds for male and female primary beneficiaries are similar – around 19-20 percent. If people claim benefits one year earlier, the benefit reduction is generally 6.7 percent.[4] But the latter effect persists for every year after benefits are claimed. Thus, it seems unlikely that extra earnings of 19-20 percent for perhaps a year or two would be nearly enough to offset the benefit reduction even if we were just considering a surviving (female) spouse with lower costs of living. There is the potential for higher benefits for some because of benefit recomputation replacing higher earnings with lower earnings, although that is likely only for certain occupations for which the earnings profile implies favorable recomputation.

[3] "CRS Summary: H.R. 149: Social Security Earnings Test Repeal Act of 2007." The Library of Congress, http://thomas.loc.gov.

[4] For cohorts born in 1938 and after, it is slightly more complicated, with a smaller reduction in benefits for benefits claimed more than three years before the FRA; see Neumark and Song (2013).

Song and Manchester (2007) find that the effect *on earnings* of removing the RET at the FRA is concentrated among high earners (for both males and female primary beneficiaries), suggesting that there may be little effect on those more likely to be poor at old ages. The elimination of the RET could still encourage earlier benefit claiming among those with lower incomes who were not claiming benefits because their earnings prior to the change exceeded the earnings test threshold, but who have little ability to increase or manipulate their earnings in response to changes in the RET. In contrast, if low earners simply are unlikely to have earnings after age 65, then we would not expect much claiming response for them.

III. Empirical Approach

The fundamental goal of the empirical analysis is to test whether the 2000 reforms eliminating the RET had adverse consequences for the incomes of much older women whose own behavior or spouse's behavior was affected by the elimination of the RET. The challenge in estimating this kind of effect is the usual one of the counterfactual. We would like to compare observed outcomes for cohorts affected by the elimination of the RET to what their outcomes would have been had the RET not changed. An inherent difficulty of this approach is that the policy change applies to everyone in the affected cohort, which makes identification less compelling than in a case with geographic policy variation so that only some members of particular cohorts are affected.

We begin by estimating reduced-form models in which we identify the effect of the elimination of the RET from intercohort changes. To keep things simple, consider first a world without couples, in which women in different cohorts face different RET rules, choose when to claim benefits in part based on these rules, and then for whom we subsequently observe Social Security benefits and other sources of income. The corresponding reduced-form equation for age at claiming, for example, is of the form:

$$AgeClaim_i^w = \alpha + \beta EET_i^w + X_i^w \gamma + \varepsilon_i \,, \tag{1}$$

where $AgeClaim^w$ is the age of the woman/wife when she claimed Social Security benefits, EET^w is a dummy variable equal to one for cohorts of women for which the RET at the FRA and above was eliminated (less than age 70 in 2000), and X^w is a vector of demographic control variables (race/ethnicity, educational attainment, and marital status). X^w also includes a dummy variable for cohorts reaching age 65 in 2003 or later, for which the FRA was greater than 65 and benefits before the FRA were lower than for earlier cohorts. These individuals would be predicted to claim benefits later and have lower benefits conditional on age (see Neumark and Song, 2013). We also estimate versions of this model in which EET^w is broken up into five dummy variables for those age 69 in 2000, age 68, age 67, age 66, and age 65 or younger. The youngest group had the highest "exposure" to the elimination of the RET, whereas the oldest group was exposed only at age 69, and hence should have changed its behavior less.

This regression identifies the effect of the elimination of the RET from differences in outcomes between cohorts that were and were not affected by the elimination of the RET. It is estimated for subsamples of women age 70 or older, all of whom should have claimed their Social Security benefits, albeit at ages that may differ because of exposure to the elimination of the RET. If we had policy variation across states then we could identify these from the states where the policy change does not occur and identify the policy effect from interactions at the ages and cohorts affected by the policy change in the states where the policy change occurs, paralleling a difference-in-differences specification. But with only national variation we have to be more restrictive and use the cross-cohort differences. Similarly, we do not have a reversion of the policy so that still-younger cohorts unaffected by the change that could be studied to help distinguish between a policy effect and underlying cross-cohort trends.[5]

Identification of the effect of the elimination of the RET in equation (1) requires the identifying assumption that there are not other sources of differences in these outcomes across cohorts. We provide evidence of this by comparing results for a wider and narrower range of birth cohorts. With the narrower range, it is less likely that other sources of differences are important. We can also learn something from the estimates of the specification in which EET^{w} is broken up into five dummy variables that capture differences in the number of years and youngest age of "exposure" to the elimination of the RET. The effect should be stronger the younger an individual was at the time of the elimination of the RET because there was more time to adjust behavior. Nonetheless, simply using cohort variation might be viewed as not entirely compelling, but is the best we can do this in this case.

We estimate this model for age at claiming as explained above, to establish a baseline for our analysis and touch base with the earlier literature. In particular, we expect those exposed to the elimination of the RET to claim benefits earlier. We are more interested, however, in the consequences for income – and more specifically the incidence of low income. We therefore also estimate a similar specification for Social Security benefits and then for income relative to poverty thresholds. We denote these variables generically, for women, as Y^{w}.

In thinking about these latter specifications, it is instructive to think about the counterfactual used in equation (1), which has implications for how to index Social Security benefits in the model for these benefits or for all income. To be concrete, consider the example

[5] By way of contrast, Krueger and Pischke's (1992) paper on the "notch babies" (born 1917-1921) who faced much lower social benefits than either immediately older *or* immediately younger cohorts, which allows the researchers to infer the effect of lower benefits on labor supply by looking for a break in the longer-term time-series behavior of labor supply for this narrow cohort (which turns out to be weak). Snyder and Evans (2006) present a similar analysis of the effect of the benefits change on mortality (and also report higher post-retirement labor supply for the affected cohorts).

of using data only on the 1931 and 1930 birth cohorts; because these cohorts are aged 69 and 70, respectively, in 2000, the first is affected by the elimination of the RET, and the second is not affected. Because of the change in the policy, members of the 1931 birth cohort would be expected, on average, to claim benefits earlier, resulting in lower benefits. The reduced-form regression uses unaffected cohorts – in this example, the 1930 birth cohort – as a counterfactual for what benefits would have been, and when they would have been claimed – for the 1931 birth cohort.

For the age at claiming, there is no inherent problem. As long as we assume there were no other cohort effects aside from the elimination of the RET that would have led to different ages at claiming across these cohorts – which is the key identifying assumption – the estimate of β identifies the cohort effect.

However, when the outcome is benefits, getting the counterfactual right is more complicated. As in the analysis of the effect on the age at claiming, we want to use outcomes for the 1930 cohort as the counterfactual. But the indexation of Social Security benefits implies that benefits for the 1930 cohort may not correctly estimate the counterfactual for the 1931 cohort had the RET not been eliminated. In particular, there is a two-step indexation process. First, for each individual, the average wage index (AWI) as of age 60 is used to bring earnings prior to age 60 up to current nominal levels in setting the primary insurance amount (PIA). In the case of just two cohorts, this means that the PIA of the older 1930 birth cohort would have to be inflated by the AWI for 1991 relative to 1990 to get the right "counterfactual PIA" for the 1931 birth cohort. This is important because otherwise the benefits of the 1931 birth cohort may appear too high – masking the reduction in benefits from claiming earlier because of the elimination of the RET. The PIA is subsequently indexed by the CPI-W, which implies that we have to use this index from age 60 to the year of observation to have comparable current dollar benefits for the two cohorts.[6]

Thus, when we look at Social Security benefits, we first multiply by the ratio of the AWI in 1995 to the AWI when the person was aged 60. The year 1995 is chosen somewhat arbitrarily, but this means that we put all workers' benefits on an equivalent footing, in terms of the AWI, to a worker who was age 65 in 2000 (the first cohort exposed to the elimination of the RET beginning at the FRA). Then, because benefits are observed at different years depending on age and when a person is observed in the HRS, we multiply this adjusted figure by the ratio of the CPI-W in 2013 to the CPI-W in 1995, to express all benefits in 2013 dollars. Benefits can still vary across individuals (and couples) based on when they claimed benefits – which affects the relationship between PIA and benefits. But by doing this indexation, we isolate the variation in benefits due to age at claiming.

For our main analysis of the effects of the RET elimination on poverty or low-income, we have to construct total income across Social Security benefits and other sources. For non-Social Security income, we simply index by the CPI-U to express all amounts in 2013 dollars (since the same counterfactual issue arising from wage indexing of benefits does not arise). Thus, when we construct total income we are indexing Social Security benefits and other income differently.

We also use an instrumental variables strategy that estimates the effect of age at claiming on the outcomes we study, using as an instrumental variable (IV) the same single or five-way indicator of exposure to the elimination of the RET above (with the former defined above as EET^{w}). Under a local average treatment effect (LATE) interpretation, this approach estimates the effects of early claiming induced by the elimination of the RET on subsequent labor market outcomes for older women. The LATE parameter is likely of particular policy interest in this

[6] For more on indexation of Social Security benefits, see http://www.ssa.gov/oact/cola/AWI.html, http://www.ssa.gov/oact/cola/Benefits.html, http://www.ssa.gov/oact/cola/latestCOLA.html, and http://www.ssa.gov/oact/COLA/colaapplic.html (viewed March 29, 2015).

setting, because it focuses on how the elimination of the RET affects income-related outcomes. However, the variation in the IV still comes from these cross-cohort differences and hence could, in principle, reflect influences aside from the elimination of the RET.

In the IV approach, the empirical strategy is to estimate a model for the age at benefit claiming as a function of controls assumed to be exogenous, as well as exposure to the elimination of the RET. This serves as a first-stage for a "structural" equation relating income and poverty-related outcomes for women.

Returning to the simple example of women only (described above), women in different cohorts face different RET rules, and choose when to claim benefits, in part, based on these rules. We then subsequently observe Social Security benefits and other sources of income. The structural equation of interest takes the form:

$$Y_i^{wa} = \delta + \theta AgeClaim_i^w + X_i^w\lambda + \xi_i \ , \qquad (2)$$

where Y^{wa} is an income-related measure defined for woman *i*, as of age *a*. $AgeClaim^w$ and X^w are defined as before.

We assume that age at claiming is affected by the same variables in *X*, as well as exposure to the elimination of the RET. The simplest form of this variable (again, in our example of women only) is a dummy variable equal to one for cohorts for which the RET was eliminated, EET_i^w, where the *w* superscript indicates that this dummy variable is based on the woman's (or wife's) birth cohort. Therefore, our first-stage regression is:

$$AgeClaim_i^w = \mu + \lambda EET_i^w + X_i\kappa + \nu_i \ . \qquad (3)$$

The model, as written, is the standard linear equation framework with constant coefficients. Under the assumptions that (a) EET^w predicts $AgeClaim^w$, and (b) that Y^a is uncorrelated with EET^w, conditional on $AgeClaim^w$ and *X*, then in this framework *EET* is a valid instrumental variable (IV) for *AgeClaim*^{w} in equation (2).[7]

In this linear equation, constant coefficient model, the IV estimate simply provides an estimate of the effect of exogenous variation in age at claiming on later outcomes. We are not interested in this general question, although if the linear, constant coefficient model held, the IV estimate of the effect of the age at claiming would be more believable than the OLS estimate, which might, for example, be biased by unobserved variation in health that is associated with both lower income and earlier claiming.

Instead, we exploit the LATE framework, which does not impose a constant effect of age at claiming on later outcomes. Under the LATE interpretation the IV estimate identifies the effect of changes in the age at claiming for those induced to claim earlier because of the elimination of the RET (Imbens and Angrist, 1994) – precisely our question of interest. This interpretation requires the additional assumption of monotonicity, which in this context means that if exposure to the elimination of the RET on average causes earlier benefit claiming, then exposure to the elimination of the RET does not cause some people to claim later. Like the exclusion restriction, this is not testable, but there is no reason to think the assumption would be violated in the case of the RET and the age at claiming.

To this point we have described models that are estimated for women only. We also, however, estimate models that capture the effects of the elimination of the RET for both women and their husbands, since both can affect benefits depending on whether the woman is claiming her own benefits ("entitled") or spousal/survivor benefits.

The reduced-form equation incorporating the exposure of both husbands and wives to the

[7] One could imagine introducing some function of year of birth and including this in both the first-stage and structural equation to allow for other sources of change by birth cohort. However, because the actual effect of the elimination of the RET is to introduce variation in years of exposure to the elimination of the RET that goes from one for those aged 69 in 2000, to five for those aged 65 or younger in 2000, it would in practice be difficult to distinguish between the effects of the elimination of the RET by cohort and other cohort effects.

elimination of the earning test is simple, and takes the form:

$$AgeClaim_i^w = \alpha + \beta EET_i^w + \beta' EET_i^h + X_i^w \gamma + \varepsilon_i \,. \qquad (4)$$

EET^h is the dummy variable indicating that the husband was younger than age 70 in 2000.[8]

For the IV approach, the structural equation now contains two age at claiming variables, as in:

$$Y_i^{wa} = \delta + \theta^w AgeClaim_i^w + \theta^h AgeClaim_i^h + X_i^w \lambda + \xi_i \,. \qquad (5)$$

And there are two first-stage equations, one for each of the age at claiming variables:

$$AgeClaim_i^w = \mu + \lambda^w EET_i^w + \lambda^h EET_i^h + X_i \kappa + \nu_i \,, \qquad (6)$$

and

$$AgeClaim_i^h = \varphi + \psi^w EET_i^w + \psi^h EET_i^h + X_i \kappa + \varsigma_i \,. \qquad (7)$$

IV. Data

To study the effects of the elimination of the retirement RET, we use the Health and Retirement Study (HRS).[9] The HRS is a longitudinal survey administered every two years, interviewing a nationally representative sample of individuals older than age 50. The HRS began collecting data in 1992 and the most recent available data in the HRS is from the 2012 survey wave.

We impose several sample restrictions on the data to obtain the final samples. We begin by limiting the HRS data to individuals whose Social Security claiming age is not missing and is between ages 62 and 71. We remove observations that report receiving annual Social Security benefits less than \$6,000 or greater than \$35,500.[10] Finally, for most of the analyses presenting

[8] We also estimated versions of this specification where we included two sets of five dummy variables – one for the woman and one for her husband – indicating one, two, three, four, or five years of exposure to the elimination of the RET. These led to qualitatively similar results to those reported in the tables that follow.

[9] We use the RAND version of the HRS.

[10] These limits are imposed based on the CPI-U inflation adjusted Social Security benefits to 2013 dollars.

our main results, we limit our data to women in two age ranges. Our “age 70+ sample” consists of women ages 70 and 71, and our “age 75+ sample” consists of women ages 75 and 76. We use a two-year window because the HRS is conducted every two years. The age of the individual is measured at the end of the reference year, typically the calendar year prior to the interview year.[11] This results in a sample of 2,974 women in the age 70+ sample and 1,958 women in the age 75+ sample. Table 1 contains the number of observations by age in the these two samples, and by whether the person’s birth cohort made them less than age 70 in 2000 and hence subject to the elimination of the RET, and older than 70. Later, we report on the robustness of our key conclusions – especially changes in the results as women age – to looking at a much broader range of ages.

We also create subsamples that include only women who can be matched to a unique husband; observations that report multiple husbands complicate matters because it is unclear which husband’s age – and hence exposure to the elimination of the RET – affects the woman’s Social Security benefits. We use these subsamples to study the effects of the husband’s claiming decision on the family’s and wife’s outcomes. To construct this sample, we removed women from the sample who never report a spouse or report multiple spouses during the HRS sample.[12] We also limit this sample to women whose husband’s Social Security claiming age is not missing and is between ages 62 and 71. The corresponding sample sizes by age are also reported in Table 1.

One of the primary outcomes of interest is whether or not the woman lives below particular low-income thresholds, such as the poverty line or a multiple of it. Unfortunately, the

[11] For the HRS income questions, the reference year is typically the calendar year prior to the interview year.
[12] In implementing this method, we require women to be observed as “married” at least once during the HRS sample. We also exclude same-sex couples because all of the data occurs before the U.S. Supreme Court’s decision in *United States v. Windsor* in 2013, which allowed same-sex married couples to claim on their spouse’s work history in the same manner as heterosexual married couples do.

HRS does not provide a measure of whether the household lives below the poverty line until the 2002 wave. For data prior to 2002, we construct a measure of poverty using the Census Bureau's definition of poverty. To determine whether the household is residing in poverty, the Census Bureau uses the number of individuals residing in the household, the total income of the members of the household, and the age of the related children residing in the household. The HRS only includes the total income of the respondent and spouse (if one is present) in its measure of household income. To obtain the age of the related children, we merge in the RAND HRS family data (which includes information on the number of children residing with the respondent from 1992 to 2008). However, in line with the earlier discussion of the correct construction of Social Security benefits for the counterfactual, we use a constructed income measure, which is whether adjusted benefits (as described earlier) plus other household income is below the poverty line or a multiple of it.[13]

Descriptive statistics are reported in Table 2. These reflect features we would expect. For example, the average age at claiming benefits (self-reported) is between 63 and 64. The widowhood rate is much higher for the all-female sample than for the women with husband's sample, regardless of age. While women for whom we observe a husband during the sample period have a lower widowhood rate than women overall (0.17 compared to 0.26), the probability of being widowed dramatically increases by age 75.

One thing to note in Table 2 is the very low incidence of poverty in the samples of women with unique husbands observed, in the last two columns (although, as expected, the

[13] In order to check how accurate our calculated poverty measure is at predicting the incidence of poverty, we apply our measure to the entire HRS sample for the 2002 to 2008 waves and compare the incidence of poverty to the HRS calculated poverty measure. (Observations with a zero weight are excluded to avoid including observations residing in nursing homes.) The calculated measure is able to accurately predict whether or not the household resides in poverty in 96 percent of cases, and whether or not the household is below 200 percent of the poverty line in 93 percent of cases. (This calculation is done without our adjustment of benefits by the AWI.) Note that our key results for older women (e.g., the age 75+ sample), are not affected by this imputation because all observations come from 2002 or later.

percentage of these women living in poverty increases between the ages of 70 and 75). This arises for two reasons. First, the adjusted Social Security benefits that we use reduce the estimated poverty rate, by inflating benefits for some. Second, this subsample of women has very low poverty rates, in part, because they were married at some point in the period covered by the HRS, and also, in large part, because if they were divorced they have not had multiple husbands in the period covered by the HRS.[14]

As a consequence of the very low share of women in this sample who are below the poverty line, we focus instead on whether women are below 150 percent or 200 percent of the poverty line. The proportions living below these thresholds are higher, as Table 2 shows, likely in part because the sample selection necessary to define the treatment of husbands is not as severe for this threshold. These thresholds are still of considerable interest because they would still be considered to be low-income thresholds. In addition, the results may differ for these higher income thresholds, since – as Figure 1 shows – those with higher earnings initially are more likely to increase labor supply in response to the elimination of the RET. That is, at these thresholds we may be somewhat less likely to see older women affected by the elimination of the RET having very low incomes; the fact that we, nonetheless, find such an effect for these above-poverty thresholds, thus strengthens our conclusions.

Our analysis uses the public HRS data, in which we cannot identify what kind of benefits

[14] The poverty rates for the women we drop to obtain our women with husbands sample at age 70 are as follows: (1) 6.7 percent live below the poverty line, (2) 19.1 percent live below 150 percent of the poverty line, and (3) 33.3 percent live below 200 percent of the poverty line. At age 75, the poverty rates for these individuals are as follows: (1) 2.3 percent live below the poverty line, (2) 13.8 percent live below 150 percent of the poverty line, and (3) 27.1 percent live below 200 percent of the poverty line.

women are claiming.[15] Ideally, we would also (or instead) use the confidential administrative Social Security records that are available for many HRS respondents, which would give us an administrative measure of the age at claiming, and also identify what kind of benefits women are claiming and thus perhaps better pin down whose exposure to the elimination of the RET (hers or her husband's) are driving benefits.[16]

However, the administrative data are complicated, and our work with these data has identified two issues pointing to problems with identifying the relevant claiming date in the administrative data. First, there are substantial discrepancies between self-reported age at claiming and the age at claiming calculated using the administrative data.[17] The number of discrepancies is large, and many are four months or larger, or even 13 months or larger. There is a difference between cohorts born before 1931 – for whom claiming dates are almost always based on recall before the first survey for the relevant cohort – and those born in 1931 or later – for whom claiming dates come from recall limited to the two years between interviews and should be more accurate.[18] The errors are larger for those born before 1931, as we might expect. But there are many large errors for the younger cohorts as well. The difference across birth cohorts is particularly problematic because birth year 1931 is the dividing line between those who are or are not affected by the elimination of the RET.

[15] Appendix Table A1 gives information on the type of benefits women earned over the 1995-2013 period. In addition to showing the percentage of women claiming on their own work histories, Table A1 displays the fraction claiming "spousal" or "survivor" benefits. Spousal benefits are equal to one-half the husband's benefit amount. When the husband dies, the wife can chose to continue her current benefits (either own or spousal) or claim survivor benefits, which are equal to the husband's benefit. Typically, the highest benefit amount is claimed. While men can claim either spousal or survivor benefits, it is typically women who receive these benefit types. Table A1 shows a substantial share of women are receiving benefits based on their own earnings histories, especially in more recent years.

[16] Even for a woman claiming her own benefits her age at claiming may be affected by her husband's decision and hence his exposure to the elimination of the RET – for example because of leisure complementarities. For work on this topic see Hurd (1990) and Coile (2004). For a more recent study using French data see Stancanelli and van Soest (2012).

[17] See Appendix Table A2.

[18] The HRS cohort was born 1931-1941 (although spouses can be born in different years), and was first interviewed in 1992. The Study of Assets and Health Dynamics (AHEAD) and Children of Depression (CODA) cohorts were born earlier, but not interviewed until later (first in 1993 for the AHEAD cohort, and 1998 for the CODA cohort). Thus, claiming dates for those born before 1931 are generally recalled.

We could, of course, assume the administrative data are correct. But this leads to the second issue. Specifically, when we try to replicate the Song and Manchester (2007) results using the HRS administrative data, we are unable to replicate their findings. Yet, when we use the HRS public data and the self-reported age at claiming, we are able to replicate the Song and Manchester (2007) findings. This is reported in Appendix Table A3.

To create a sample similar to Song and Manchester (2007), we limit the administrative data to men and women who are observed claiming benefits and who have not claimed disability benefits. In the public data, we limit the sample to individuals who report a claiming age of age 62 or older to avoid including individuals who are receiving disability benefits. Song and Manchester (2007) conduct the analysis for two treatment groups. One of the treatment groups includes those who had attained ages 65 to 69 by January 1 of the calendar year. The second treatment group includes those who turn age 65 during the year. We focus on replicating the estimates for the first group because the RET is completely eliminated for these individuals. While the estimates for those turning age 65 during the calendar year are informative, these individuals face a prorated RET during the months prior to turning age 65.[19] In contrast, the earnings test is completely eliminated for those who have reached age 65 prior to January 1 of the calendar year.

Appendix Table A3 provides our estimates using the HRS administrative and public sample. Using the HRS public we find statistically significant increases in claiming for the treatment group that are in line with the Song and Manchester (2007) findings. Using the HRS administrative data, however, we find small and statistically insignificant changes in claiming behavior for the treatment group. Together, these findings lead us to proceed using the public HRS data.

[19] Note this assumes that the FRA is age 65. The same is true for FRAs above age 65. In the year the individual reaches the FRA, the individual faces a prorated earnings test during the months prior to turning the FRA.

V. Results

Individual Benefits

We begin with models estimated for Social Security benefits and the age at claiming to test for the expected effects (in some cases reflected in earlier literature as well) that the RET leads to earlier claiming and lower benefits. In Table 3, we report OLS estimates, for women in isolation, of regressions for the age at claiming benefits. The specifications vary in two ways: First, whether we include a single dummy variable for "exposure" to the elimination of the RET (less than age 70 in 2000), or instead five dummy variables that capture the age at which the elimination of the RET occurred; and second, whether we use the full sample, or a more restricted range of birth cohorts. As columns (1) and (3) of Table 3 show, for both samples we find that the elimination of the RET reduces the age at claiming by eight to nine months – estimates that are statistically significant at the one-percent level. In columns (2) and (4), where we instead use the richer characterization of exposure to elimination of the RET, we find, for the most part, that the younger the woman was when the RET was eliminated, the earlier she claimed. The estimates range from 6.5 to 7.8 months for those who were 69 when the RET was eliminated, to about 8.2 to 9.6 months for those who were age 65. The relationship is monotonic with one exception in each of these two columns; as noted earlier, this match between the estimated effects of years of exposure to the elimination of the RET and the effects we would predict suggests that we are not picking up spurious cohort effects unrelated to the elimination of the RET. We do note that effects of eight months or so seem large, if only those who have not claimed benefits by age 65 are affected by the elimination of the RET; it is possible, though, that some claiming prior to age 65 is also affected, perhaps because of the forward-looking nature of joint decisions regarding claiming and labor supply.

In Table 4, we report OLS estimates of regressions for the level of benefits. The estimates reflect what we would expect based on the age at claiming. In columns (1) and (3), we see that benefits (annual) are lower by about $650-$800 for cohorts for whom the RET was eliminated. In columns (2) and (4) we see that benefits are to some extent successively lower the younger the age at which one was exposed to the elimination of the RET (the big exception is for some anomalous results for those age 66 in 2000). Most notably, the estimates for the cohorts exposed at the youngest ages are largest and more strongly statistically significant. Of course, we get the most observations on those who were 65 or younger in 2000, which is reflected in considerably lower standard errors.

Table 5 reports the IV (two-stage least squares) estimates of the models for benefits, where the first stage is the corresponding age at claiming regression in Table 3. In these models, the age at claiming is treated as endogenous. Exposure to a higher FRA (for those who become 65 in 2003 or later) is included in the model to control for other sources of changes in age at claiming and hence benefits. Exposure to an FRA greater than age 65 is not used as an IV for claiming for two reasons. First, the changes in the FRA were accompanied by reductions in benefits at each age for claiming before the FRA, and hence this variable cannot be excluded from the structural equation. Second, we want to isolate the effect of the elimination of the RET in the local average treatment effect (LATE) as well. In columns (1) and (3) we use the single dummy variable for exposure to the elimination of the RET as an instrumental variable, and in columns (2) and (4) we use the five dummy variables capturing years of exposure to the elimination of the RET as instrumental variables (in both cases matching the first-stage estimates in the same columns of Table 3).

In all four columns of Table 5, the estimates are in close correspondence, with benefits that are lower by about $80-$90 for each month earlier a woman claims owing to the elimination of the RET (using a LATE interpretation of the estimate). Note that the estimates here are the opposite sign, because we estimate the effect of the age at claiming, which occurs *earlier* for those affected by the elimination of the RET. The estimates in all four columns are statistically significant at the one-percent level. The table also reports the F-statistics from the first-stage regressions (reported in Table 3). These are always very large – in the 80-100 range when using just one IV, and the 16-22 range when using the five dummy variables. Note also that benefits are lower for those exposed to a higher FRA. Although they claimed later on average (Table 3), benefits were reduced and, hence, should be lower conditional on age at claiming. Thus, the results for women in isolation indicate that the elimination of the RET resulted in earlier claiming of Social Security benefits – by about 2/3 of a year – and lower benefits, by a bit less than $100 per month earlier that benefits were claimed. The magnitudes of the benefit effects seem reasonable. Given that benefits are reduced about 0.6 percent for each month earlier that one claims, and annual benefits (Table 2) average $15,547, the direct implication of claiming one month earlier would be an annual benefit reduction of $93. But recall that those likely to claim earlier owing to the elimination of the RET are higher earners with higher benefits.

Family Benefits

Having established these results for women in isolation, we next turn to the effects of the elimination of the RET on poverty or low-income status of older women. In this case, we estimate specifications for family Social Security benefits. The approach parallels the above analysis for women only, except that we now study a sample of women with one husband observed during the period covered by our HRS data so that we can determine the birth cohort of the husband that would have determined *his* exposure to the elimination of the RET. In addition, we estimate models including variables capturing both the wife’s and the husband’s exposure to the elimination of the RET.[20]

We begin in Table 6 with the estimates of models for the age at claiming Social Security benefits. For both cohort ranges, we report estimates for both women and their husbands. We also estimated these models using the five dummy variables for the years of exposure to the elimination of the RET – for both husbands and wives (a total of 10 dummy variables in each model). In the interest of space we do not report these specifications, even though they serve as the first stages for estimates reported below. For the larger cohort range, column (1) reports estimates for wives' ages at claiming, and column (1') for their husbands.[21] We see, not surprisingly, that the wife's age at claiming is affected by whether she was less than age 70 in 2000, and the husband's age at claiming by his age in 2000. Both estimates are negative and statistically significant at the one-percent level, indicating earlier claiming among those for whom the RET was eliminated. There is no evidence of "cross" effects between spouses. In columns (2) and (2') we restrict the cohort range, and find similar results.

Next, Table 7 reports OLS estimates of the equations for family Social Security benefits, taking account of both husbands' and wives' exposure to the elimination of the RET. However, recall our earlier caveat that these results are based on a highly selected sample of women with a unique husband observed in the HRS. We find fairly strong evidence that when husbands were exposed to the elimination of the RET, family benefits are lower. Conditional on the exposure of husbands, the effects are the opposite sign for women's own exposure. This finding is a bit puzzling, and may reflect other influences on benefit levels associated with whether the wife was less than age 70 in 2000, conditional on whether the husband also was less than age 70 in 2000.

[20] Note that the subsample of women we use is smaller because of this restriction. In results not reported in this paper, we verified that the results for this subsample, but looking only at women's benefits, and estimating the same specifications as in Tables 6-8, qualitatively replicated the results for the larger sample of women covered in those tables, although the estimated effects on benefits were a bit smaller and not always statistically significant.

[21] We label these columns (1) and (1') because these *two* equations serve as the first stages for the model estimated below treating age at claiming of both husbands and wives as endogenous.

Another way to think about the estimates, though, is that both spouses are likely to be affected similarly by the policy given small age differences between most husbands and wives. Alternatively, the "experiment" of thinking about the partial effect of eliminating the RET only for wives or only for husbands is of limited interest. In this case, the fact that the effects for husbands are larger indicates that, on net, family benefits are likely to be lowered somewhat because of the elimination of the RET. Indeed this is shown in the penultimate row of the table, where we report the combined effect of eliminating the RET for both husbands and wives, leading to significantly lower annual benefits of around $1,500-$1,600.

Table 8 turns to estimates of models where we treat the age of claiming benefits of both wives and husbands as endogenous. Here, we restrict attention to the "single" treatment variables regarding exposure to the elimination of the RET as opposed to treatment variables defined by single-year age groups, which would lead to 10 instrumental variables (although we verified that results were qualitatively similar). The specification has two endogenous variables (wives' and husbands' age at claiming), and two instrumental variables (wife less than age 70 in 2000, and husband less than age 70 in 2000). The estimates indicate that family benefits are lower by about $475 a month when husbands claim benefits earlier because of exposure to the elimination of the RET; these estimates are statistically significant at the one-percent level. Paralleling the OLS regression estimate for benefits, the sign is the opposite for the age of claiming of wives, suggesting that husbands' behavior offsets this is some way. However, as we argued before, the effect of most interest is the combined effect for husbands and wives. The estimates of the combined effect indicate that family benefits are lower by about $200 a month for each month earlier that benefits are claimed as a result of the elimination of the RET.

The Incidence of Low Income Relative to Poverty

Finally, we turn to estimated effects on the incidence of low income as defined by multiples of the poverty threshold. The structure of the tables is exactly the same as Tables 7 and 8. Note that the sample is slightly smaller here because we use the Census Bureau definition of poverty, which excludes those residing in institutionalized facilities (e.g., nursing homes) from the poverty calculation.[22] Given the small differences (seven observations) we do not report the first-stage estimates for the poverty observations sample; these would be virtually the same as in Table 7. In addition, for these specifications we do not report any specifications using years of exposure to the elimination of the RET, since we did not find as clear evidence of a systematic pattern; but the IV estimates using these richer first stages were qualitatively similar.

The top panels of Tables 9 and 10 report results for the same sample restriction we used in the preceding tables – the first observations on women at least 70 years old. The bottom panels report estimates of the same models, but for an older sample of women aged at least 75 – an age at which more women are likely to depend more exclusively on Social Security benefits, and hence for whom we might be more apt to see the emergence of adverse effects of the elimination of the RET on family income. For the sample of women for whom we observe a husband, Social Security benefits make up 54.5 percent of total household income at age 70 and 61.4 percent of total household income at age 75.

In the top panel of Table 9, columns (1) and (2) report the estimates for the probability that income is below 150 percent of the poverty line, and columns (3) and (4) for the probability that income is below 200 percent of the poverty line. In columns (1) and (2), the point estimates for exposure of women to the elimination of the RET indicate a lower probability of living below 150 percent of the poverty line; the estimates indicate that this probability is lower by about 4.5-5 percentage points, a statistically significant difference. The estimates in columns (3) and (4),

[22] The RAND HRS does provide an indicator for whether those living in institutionalized facilities are living in poverty. We view this measure as less salient for policy analysis regarding income support provided by Social Security. Individuals residing in institutionalized facilities are likely worse off due to considerations other than income and may be required to spend a large amount of resources on care.

using the 200 percent threshold, tell a similar story, indicating slightly larger reductions of about 5.5 percentage points.

Interestingly, in the bottom panel that turns to the older sample of women first observed at age 75 or older, the results are different. Women's exposure to the elimination of the RET is now associated with a *higher* probability of living below 200 percent of the poverty line – by 3.4 to 4.5 percentage points, statistically significant in both cases. The estimates for the 150 percent threshold are of the same sign, although smaller (about 1-1.9 percentage points) and statistically insignificant. Table 10 reports the IV estimates, which tell a similar story. The differences in the results for older sample relative to the sample first observed at age 70 suggest that as women move into their mid-70s, the effect of lower Social Security benefits from early claiming comes to dominate the effects of higher earnings (and whatever effect those higher earnings had on income from savings and income from those savings).

In Tables 11 and 12 we expand the analysis to include the effects of husbands' exposure to the elimination of the RET as well. In these analyses, the evidence on the separate effects of women's exposure to the elimination of the RET, or their husbands' exposure, is weaker. Earlier, we argued that the combined effects are of greatest interest, and the lower panels of both Tables 11 and 12 indicate that the combined effect of eliminating the RET is a greater likelihood of the older women (in the 75 or older sample) living below 200 percent of the poverty line. In both tables, this effect is statistically significant at the 10-percent level. In Table 11, the estimates indicate that the probability of these older women living below this threshold is higher by about 3.5 percentage points. The point estimates for living below 150 percent of the poverty line are also positive (about 1.2-1.4 percentage points), but not statistically significant. And the estimates in Table 12 point to earlier claiming of benefits owing to the elimination of the RET increasing the likelihood of being below 200 percent of the poverty line (the opposite sign of the reported effects of age at claiming). Moreover, the effects on older women are generated from the individual effect of eliminating the RET for women. Table 11, for example, shows that this effect is approximately 2.9-3.8 percentage points – and is statistically significant using the narrow range of birth cohorts in column (4), whereas the effects of the elimination of the RET for husbands is very small.

Again, the differences relative to the results for the sample first observed at age 70 or older suggest that, as women age into their mid-70s, the effect of lower Social Security benefits from early claiming comes to dominate the effects of higher earnings (and whatever effect those higher earnings had on income from saving).[23]

We have interpreted the differences between the age 70+ and age 75+ samples as indicating that the effect of the elimination of the RET in generating low family incomes emerges at older ages for women, when the effects of lower benefits are more likely to outweigh additional labor income. To gauge whether this is a plausible interpretation, rather than simply a difference that emerges for these two specific age thresholds, Figures 2 and 3 report results for a much wider range of age cutoffs. Figure 2 reports estimates using all women, and Figure 3 reports estimates for the sample of women with husbands observed.

In these graphs, the estimates for ages 70-71 and 75-76 correspond to those reported in the tables (columns (2) and (4) of Tables 9 and 11). The other estimates come from estimating the exact same specifications, but for the age ranges shown in the figures. What we see in the figures is that, in general, the rising likelihood of having income below either the 150-percent or

[23] The estimates of models for the probability that income is below 150 or 200 percent of the poverty line reported thus far are all based on linear probability models. Especially because the incidence of these low incomes is relatively uncommon (see Table 2), it is of interest to see if results are robust to using probit and instrumental variables probit estimation. To do this, we first specified models with more parsimonious controls (to avoid perfect predictions in the probit estimation (broader education and marital status controls), and verified that results were unchanged. We then estimated all of the models in Tables 9-12 using probit or instrumental variables probit. The estimated effects, and hence the qualitative conclusions, were very similar. These results are available from the authors upon request.

200-percent threshold (relative to poverty) is rising as women age; it does not simply occur for the comparison between the 70-71 and 75-76 year-old samples. This fairly persistent increase is particularly apparent in Figure 2, which is based on a more representative sample. It is also especially apparent for the 200-percent threshold for which it occurs throughout the age range shown, whereas for the 150-percent threshold it is most apparent over the earlier part of the age range shown and then levels off. We interpret these figures as solidifying the interpretation of our results that as women reach quite old ages, adverse effects of the elimination of the RET on their incomes emerge.

VI. Conclusion

The elimination of the RET in 2000 for those between the FRA and age 69 was intended to boost employment of those in this age range who might be deterred from working because of the reduction in Social Security benefits owing to the RET. At the same time, the elimination of the RET makes those who are working more likely to claim early. Thus, this policy change is expected to increase earnings in the short-run, but reduce Social Security benefits in the longer-run, and possibly also influence the time-path of retirement assets, saving, and income from those assets. The change will have uncertain effects on family income, including benefits, at older ages – perhaps in particular for older women who are likely to outlive their spouses and reach the point where any extra labor income from the positive work incentives of eliminating the RET are no longer evident. Finding out whether this change in the RET increased the incidence of low income for older women is important not only for understanding the effects of the 2000 policy change, but also the potential effects of additional efforts to encourage work by eliminating or reducing the RET between age 62 and the FRA, when RET currently reduces benefits by $1 for every $2 of earnings.

Our findings support a few conclusions. First, we confirm past findings that the elimination of the RET led to earlier claiming of benefits for women. Second, coupling this evidence with other evidence that men also claim benefits earlier, Social Security benefits at the individual and family level are generally lower as a result of the elimination of the RET. Third, the results for poverty and low-income status tend to fit the conjecture about higher income and hence of lower incidence of low income initially – when women are at or just above age 70 – but a higher incidence of low income as women get into their mid-70s and beyond. These findings suggest that the incidence of low income among old women was increased by the elimination of the RET.

This type of analysis could be extended in ways that may yield additional insight, including: examination of labor income, saving, and other sources of income that might be affected by the elimination of the RET; identification of the role of widowhood, which increases substantially after age 70;[24] and identification of individuals and families more likely to be affected in one direction or the other by the elimination of the RET, which may provide more useful information on how those between 62 and the FRA who are lower-skilled and have lower past earnings might be effected by the elimination or reduction of the RET.

[24] The percentages widowed in our "married" sample are 16.6 percent at age 70 and 26.5 percent at age 75 (see Table 2). For the "all women" samples, the percentages are 25.5 percent at age 70 and 38.2 percent at age 75.

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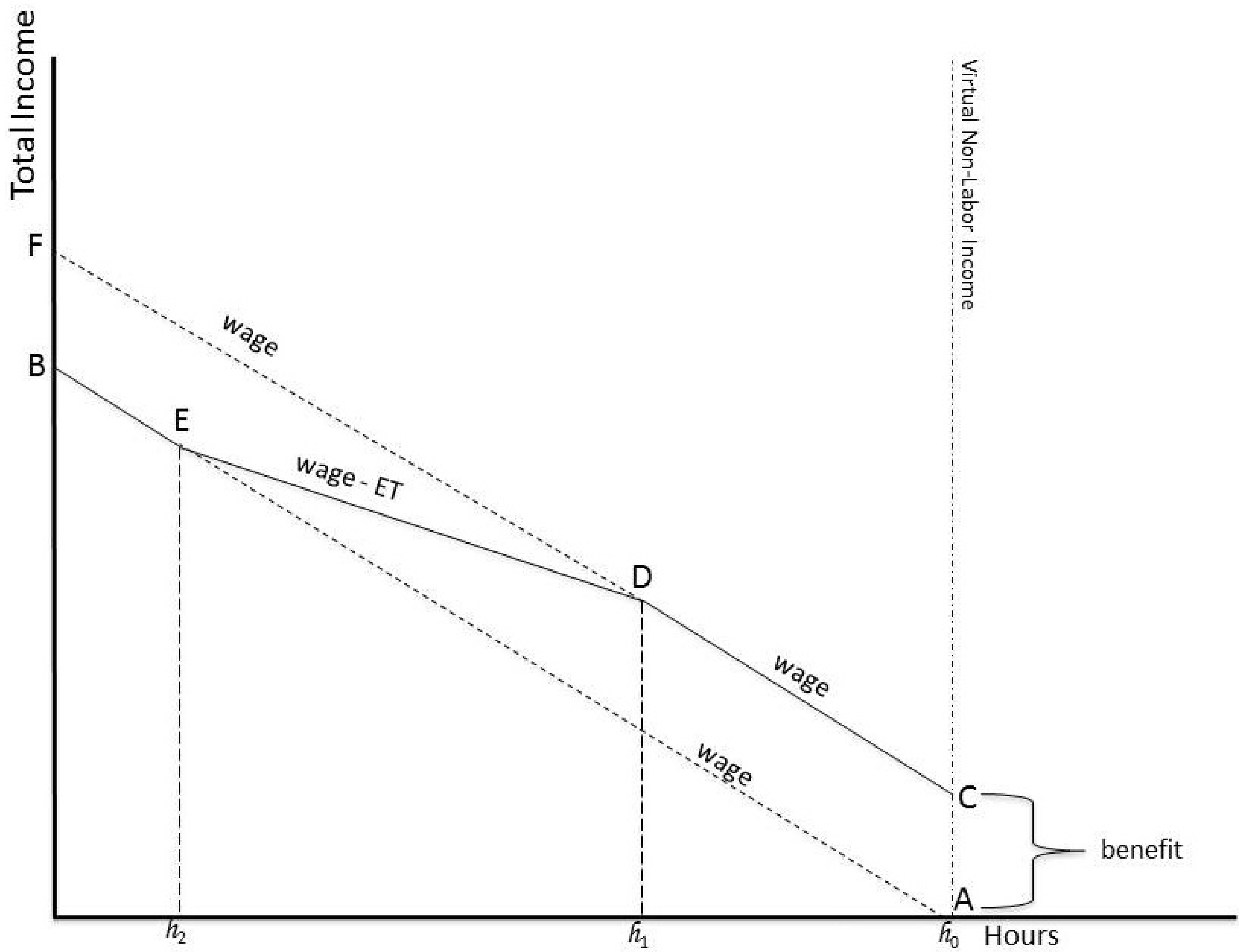
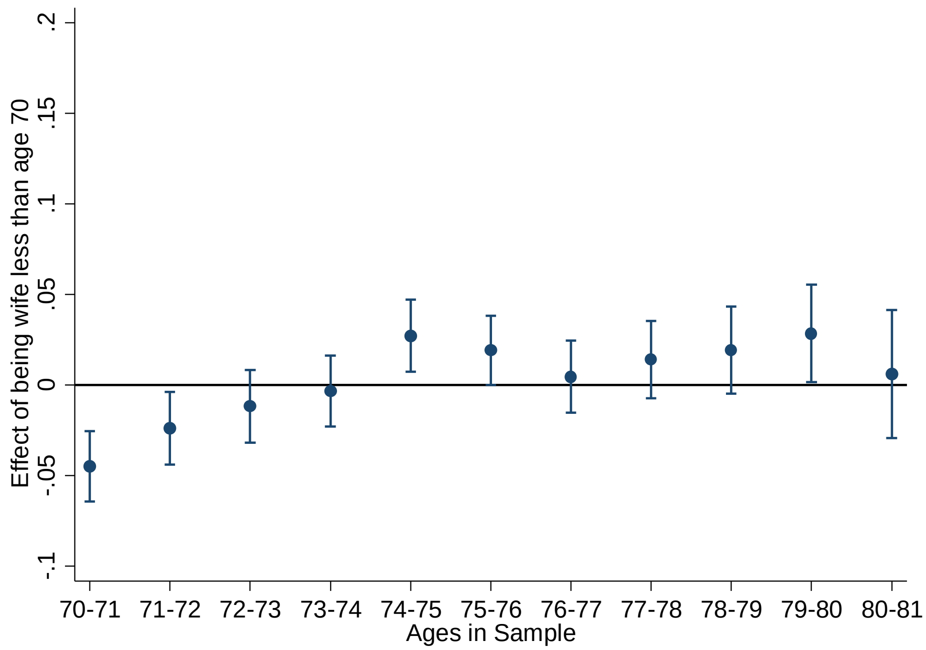
Figure 1: Budget Constraint



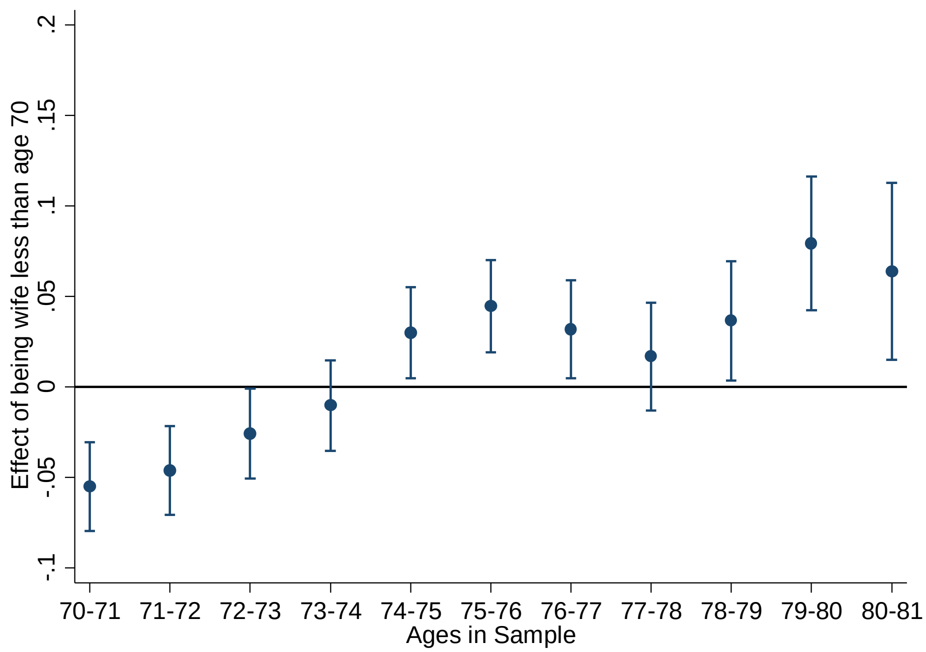
Notes: Prior to the 2000 removal of the retirement earnings test (RET), all Social Security beneficiaries younger than age 70 faced a budget constraint of CDEB. If beneficiaries younger than age 70 chose to work past the threshold (point D), these beneficiaries received a reduction in benefits (ET). After the 2000 elimination of the RET, beneficiaries who had attained the Full Retirement Age (FRA) faced a budget constraint of CDF (ignoring other taxes and benefits). Beneficiaries age 62 to the FRA continue to face budget constraint CDEB even after the 2000 changes.

Figure 2: Results on the Incidence of Low Income Relative to Poverty, Additional Ages, All Women

A. 150 Percent of the Poverty Line



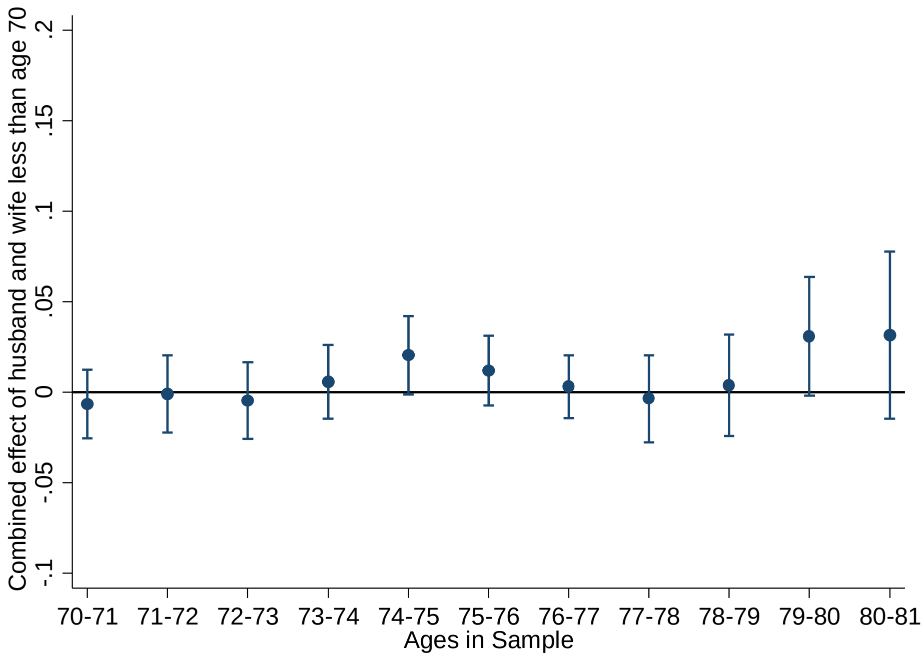
B. 200 Percent of the Poverty Line



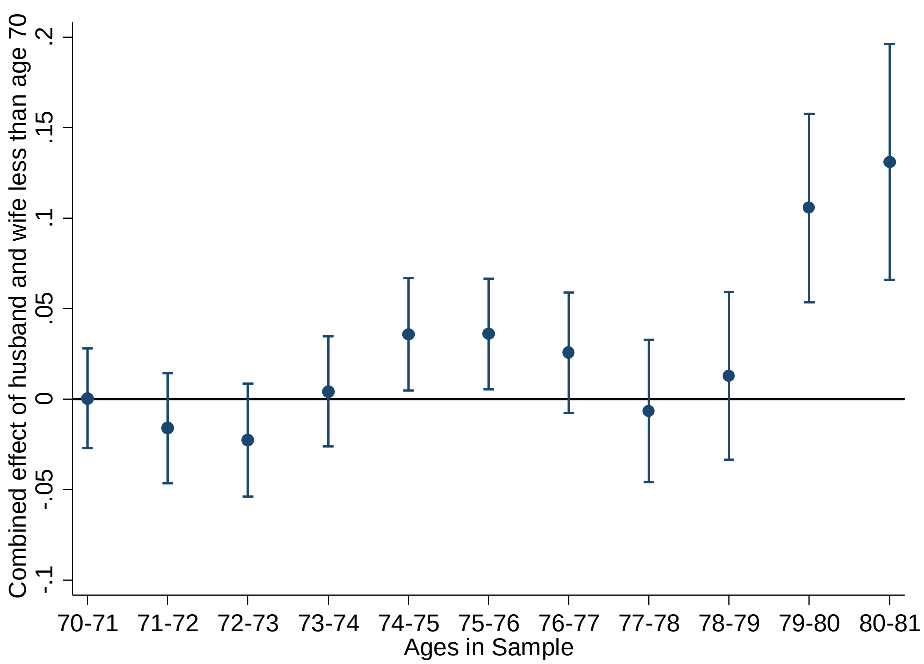
Notes: 90 percent confidence intervals are shown. Only results for those born between 1925 and 1940 are shown. The results graphed for ages 70-71 and 75-76 correspond to the estimates labeled “Less than age 70 in 2000” in Panels A and B, columns (2) and (4), of Table 9.

Figure 3: Results on the Incidence of Low Income Relative to Poverty, Additional Ages, Women with Husbands

A. 150 Percent of the Poverty Line



B. 200 Percent of the Poverty Line



Notes: 90 percent confidence intervals are shown. Only results for those born between 1925 and 1940 are shown. The results graphed for ages 70-71 and 75-76 correspond to the estimates labeled "Combined effect of husband and wife less than age 70 in 2000" in Panels A and B, columns (2) and (4), of Table 11.

Table 1: Number of Observations in Each Sample by Age and Birth Cohort

	Women				Women with husbands			
	Age 70+ sample		Age 75+ sample		Age 70+ sample		Age 75+ sample	
Age	Less than age 70 in 2000	Age 70 or older in 2000	Less than age 70 in 2000	Age 70 or older in 2000	Less than age 70 in 2000	Age 70 or older in 2000	Less than age 70 in 2000	Age 70 or older in 2000
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
70	336	1,187			166	661		
71	509	942			248	564		
75			543	542			256	345
76			406	467			188	274
Total	845	2,129	949	1,009	414	1,225	444	619

Notes: The sample is limited to individuals who are currently claiming Social Security benefits, who report Social Security benefits of greater than $6,000 but less than $35,500 in 2013 dollars, and who report an age at claiming of between ages of 62 and 71. For the "Women with Husbands" sample, we further limit the sample by removing women who never report a spouse or report multiple spouses during the HRS sample, and requiring that husband's Social Security claiming age is not missing and is between ages 62 and 71.

Source: Data from the RAND Health and Retirement Study from the 1992 to 2012 waves.

Table 2: Descriptive Statistics for Different Samples

	All Women		Women with husbands observed	
Variable	Age 70+ sample	Age 75+ sample	Age 70+ sample	Age 75+ sample
	(1)	(2)	(3)	(4)
Individual annual Social Security benefits, adjusted	15,547	19,142	14,480	17,885
Family annual Social Security benefits, adjusted	…	…	33,259	37,419
Share below poverty line	0.035	0.013	0.009	0.005
Share below 150% of poverty line	0.105	0.077	0.035	0.026
Share below 200% of poverty line	0.194	0.165	0.081	0.077
Family income excluding Social Security benefits	…	…	48,772	40,125
Social Security benefits (adjusted) as share of family income	…	…	0.595	0.686
Age	70.94	75.89	70.95	75.88
Age at claiming	63.54	63.67	63.39	63.45
< High school	0.196	0.207	0.154	0.163
High school grad or GED	0.428	0.423	0.464	0.451
Some college	0.225	0.228	0.228	0.234
College degree (BA) or higher	0.151	0.142	0.155	0.152
White	0.845	0.854	0.906	0.908
Black	0.123	0.115	0.066	0.068
Other	0.033	0.031	0.028	0.025
Current marital status: married	0.556	0.468	0.818	0.721
Current marital status: partnered	0.016	0.012	0.004	0.006
Current marital status: widowed	0.255	0.382	0.166	0.265
Current marital status: divorced	0.143	0.110	0.012	0.008
Number of observations	2,974	1,958	1,639	1,063

Notes: See notes to Table 1. Reported Social Security benefits are multiplied by the ratio of the Average Wage Index (AWI) in 1995 to the AWI when the person was aged 60, and by the ratio of the CPI-W in 2013 to the CPI-W in 1995. Other sources of income are adjusted to 2013 dollars using the CPI-U. The samples for "Women with Husbands" are slightly larger than those reported below for the regressions on poverty because observations that reside in a nursing home are assigned a missing value for the poverty variable.

Table 3: The Effect of the 2000 Elimination of the Retirement Earnings Test on the Age at Claiming Benefits in Months, Women First Observed Age 70 or Older, OLS Estimates

	Born 1918-1942	Born 1918-1942	Born 1925-1940	Born 1925-1940
Dependent variable:	Age at claiming in months	Age at claiming in months	Age at claiming in months	Age at claiming in months
	(1)	(2)	(3)	(4)
Less than age 70 in 2000	-9.03***		-7.72***	
	(0.87)		(0.87)	
Aged 69 in 2000		-7.83***		-6.54***
		(1.56)		(1.53)
Aged 68 in 2000		-7.96***		-6.67***
		(1.57)		(1.55)
Aged 67 in 2000		-10.03***		-8.73***
		(1.61)		(1.59)
Aged 66 in 2000		-8.83***		-7.53***
		(1.59)		(1.57)
Aged 65 or younger in 2000		-9.55***		-8.23***
		(1.07)		(1.06)
FRA greater than age 65	1.80**	2.32**	2.43**	2.93**
	(0.91)	(1.10)	(0.99)	(1.16)
Number of observations	2,974	2,974	2,742	2,742

Notes: See notes to Tables 1 and 2. Asterisks denote levels of significance: 1% (***) level of significance; 5% (**) level of significance; and 10% (*) level of significance. The outcome variable is the age in months that the individual began claiming Social Security benefits. The sample includes the first observation of the individual at ages 70 or 71. Only female observations are included in the sample. The specification also includes dummy variables for education (high school or GED, some college, college degree (BA) and above), race (black, white), and marital status (married, partnered, widowed).

Table 4: The Effect of the 2000 Elimination of the Retirement Earnings Test on Women's Benefits, Women First Observed Age 70 or Older, OLS Estimates

	Born 1918-1942	Born 1918-1942	Born 1925-1940	Born 1925-1940
Dependent variable:	Annual benefits	Annual benefits	Annual benefits	Annual benefits
	(1)	(2)	(3)	(4)
Less than age 70 in 2000	-807.68***		-642.46***	
	(246.89)		(246.87)	
Aged 69 in 2000		-803.03*		-645.20
		(439.92)		(435.72)
Aged 68 in 2000		-546.91		-388.36
		(444.36)		(440.12)
Aged 67 in 2000		-698.78		-532.52
		(455.28)		(450.93)
Aged 66 in 2000		6.94		162.88
		(449.12)		(444.79)
Aged 65 or younger in 2000		-1,192.34***		-1,019.94***
		(301.51)		(300.15)
FRA greater than age 65	-124.62	259.77	103.87	480.68
	(257.93)	(310.61)	(282.31)	(329.78)
Number of observations	2,974	2,974	2,742	2,742

Notes: See notes to Tables 1, 2, and 3.

Table 5: The Effect of the 2000 Elimination of the Retirement Earnings Test on Women's Benefits, Women First Observed Age 70 or Older, Two-Stage Least Squares Estimates

	Born 1918-1942	Born 1918-1942	Born 1925-1940	Born 1925-1940
Dependent variable:	Annual benefits	Annual benefits	Annual benefits	Annual benefits
	(1)	(2)	(3)	(4)
Age at claiming benefits in months (endogenous)	89.43***	91.90***	83.17***	86.52***
	(26.84)	(26.60)	(31.43)	(31.04)
FRA greater than age 65	-285.98	-282.15	-98.62	-97.38
	(240.31)	(240.47)	(263.56)	(263.88)
Instruments	Less than Age 70 in 2000	Aged 69, 68, 67, 66, 65 or younger in 2000	Less than age 70 in 2000	Aged 69, 68, 67, 66, 65 or younger in 2000
F-statistic on instrument(s)	106.9	21.80	79.07	16.25
Number of observations	2,974	2,974	2,742	2,742

Notes: See notes to Tables 1, 2, and 3. First-stage estimates are in the matching columns of Table 3.

Table 6: The Effect of the 2000 Elimination of the Retirement Earnings Test on the Age at Claiming Benefits in Months, Women with Husband Observed, Women First Observed Age 70 or Older, OLS Estimates

	Born 1918-1942	Born 1918-1942	Born 1925-1940	Born 1925-1940
Dependent variable:	Age at claiming in months	Husband's age at claiming in months	Age at claiming in months	Husband's age at claiming in months
	(1)	(1')	(2)	(2')
Less than age 70 in 2000	-11.50***	-0.98	-10.31***	-0.11
	(1.34)	(1.60)	(1.35)	(1.62)
Husband less than age 70 in 2000	1.65	-8.11***	1.79	-7.92***
	(1.22)	(1.45)	(1.22)	(1.46)
FRA greater than age 65	1.66	2.37	2.21	1.52
	(1.42)	(1.69)	(1.47)	(1.77)
Husband's FRA greater than age 65	-1.08	-2.79	0.03	-2.16
	(1.68)	(2.00)	(1.83)	(2.20)
Number of observations	1,639	1,639	1,530	1,530

Notes: See notes to Tables 1, 2, and 3. The sample is further limited to women whose husband's age at claiming age is not missing and is between the ages of 62 and 71, and who report a unique husband. Those observations that report multiple husbands complicate matters because it is unclear which husband's age affects the woman's benefit. Women who only report a spouse when "partnered" are also excluded from the sample.

Table 7: The Effect of the 2000 Elimination of the Retirement Earnings Test on Combined Family Benefits, Women with Husband Observed, Women First Observed Age 70 or Older, OLS Estimates

	Born 1918-1942	Born 1925-1940
Dependent variable:	Annual family benefits	Annual family benefits
	(1)	(2)
Less than age 70 in 2000	2,683.06***	2,740.26***
	(651.52)	(656.54)
Husband less than age 70 in 2000	-4,307.00***	-4,243.99***
	(589.17)	(593.42)
FRA greater than age 65	2,660.43***	2,952.90***
	(687.33)	(717.80)
Husband's FRA greater than age 65	-4,866.47***	-4,997.53***
	(814.32)	(892.09)
Combined effect of husband and wife less than age 70 in 2000	-1,624.34**	-1,503.73**
	(655.97)	(655.68)
Number of observations	1,639	1,530

Notes: See notes to Tables 1, 2, 3, and 6.

Table 8: The Effect of the 2000 Elimination of the Retirement Earnings Test on Family-Level Benefits, Women with Husband Observed, Women First Observed Age 70 or Older, Two-Stage Least Squares Estimates

	Born 1918-1942	Born 1925-1940
Dependent variable:	Annual family benefits	Annual family benefits
	(1)	(2)
Age at claiming benefits in months (endogenous)	-273.78***	-270.95***
	(74.33)	(80.42)
Husband's age at claiming benefits in months (endogenous)	475.14***	474.79***
	(85.06)	(87.45)
FRA greater than age 65	1,990.32**	2,830.52***
	(832.97)	(873.87)
Husband's FRA greater than age 65	-3,837.11***	-3,962.19***
	(1,092.67)	(1,170.55)
Combined effect of husband and wife less than age 70 in 2000	201.37**	203.85**
	(88.40)	(100.03)
Instruments	Less than age 70 in 2000, husband less than age 70 in 2000	
Minimum eigenvalue statistic	21.06 (7.03)	18.63 (7.03)
Number of observations	1,639	1,530

Notes: See notes to Tables 1, 2, 3, and 6. The first-stage estimates are the same as those underlying Table 6. For a discussion of minimum eigenvalue statistics and critical values, used to test for weak instruments when there are multiple instruments, see Stock and Yogo (2005).

Table 9: The Effect of the 2000 Elimination of the Retirement Earnings Test on Poverty, Women First Observed Age 70 or Older and 75 or Older Samples, OLS Estimates

	Born 1918-1942	Born 1925-1940	Born 1918-1942	Born 1925-1940
Dependent variable:	Living below 150% of poverty line	Living below 150% of poverty line	Living below 200% of poverty line	Living below 200% of poverty line
	(1)	(2)	(3)	(4)
A. Women first observed age 70 or older				
Less than age 70 in 2000	-0.0492***	-0.0450***	-0.0567***	-0.0552***
	(0.0121)	(0.0118)	(0.0150)	(0.0149)
FRA greater than age 65	0.0281**	-0.0018	0.0670***	0.0378**
	(0.0126)	(0.0135)	(0.0156)	(0.0170)
Number of observations	2,943	2,721	2,943	2,721
B. Women first observed age 75 or older				
Less than age 70 in 2000	0.0101	0.0190	0.0336**	0.0447***
	(0.0114)	(0.0116)	(0.0150)	(0.0155)
Number of observations	1,924	1,781	1,924	1,781

Notes: See notes to Tables 1, 2, 3, and 6. Poverty is based on the sum of adjusted Social Security benefits and other income relative to the poverty line (all in 2013 dollars). Note that relative to the corresponding earlier tables, we lose a small number of observations because they reside in a nursing home and are assigned a missing value for poverty. All women in the age 75 sample (Panel B) have an FRA equal to age 65. As a result, the "FRA greater than age 65" variable is excluded.

Table 10: The Effect of the 2000 Elimination of the Retirement Earnings Test on Poverty, Women First Observed Age 70 or Older and 75 or Older Samples, Two-Stage Least Squares Estimates

	Born 1918-1942	Born 1925-1940	Born 1918-1942	Born 1925-1940
Dependent variable:	Living below 150% of poverty line	Living below 150% of poverty line	Living below 200% of poverty line	Living below 200% of poverty line
	(1)	(2)	(3)	(4)
A. Women first observed age 70 or older				
Age at claiming benefits in months (endogenous)	0.0056***	0.0058***	0.0065***	0.0072***
	(0.0015)	(0.0017)	(0.0019)	(0.0021)
FRA greater than age 65	0.0182	-0.0158	0.0556***	0.0206
	(0.0129)	(0.0139)	(0.0161)	(0.0178)
Instrument	Less than age 70 in 2000			
F-statistics on instruments	99.39	78.10	99.39	78.10
Number of observations	2,943	2,721	2,943	2,721
B. Women first observed age 75 or older				
Age at claiming benefits in months (endogenous)	-0.0011	-0.0025	-0.0037**	-0.0059***
	(0.0012)	(0.0016)	(0.0017)	(0.0022)
Instrument	Less than age 70 in 2000			
F-statistics on instrument	87.23	60.69	87.23	60.69
Number of observations	1,924	1,781	1,924	1,781

Notes: See notes to Tables 1, 2, 3, 6, and 9.

Table 11: The Effect of the 2000 Elimination of the Retirement Earnings Test on Poverty, Women with Husband Observed, Women First Observed Age 70 or Older and 75 or Older Samples, OLS Estimates

	Born 1918-1942	Born 1925-1940	Born 1918-1942	Born 1925-1940
Dependent variable:	Living below 150% of poverty line	Living below 150% of poverty line	Living below 200% of poverty line	Living below 200% of poverty line
	(1)	(2)	(3)	(4)
A. Women first observed age 70 or older				
Less than age 70 in 2000	-0.0099	-0.0097	-0.0205	-0.0174
	(0.0120)	(0.0116)	(0.0174)	(0.0167)
Husband less than age 70 in 2000	0.0005	0.0032	0.0176	0.0179
	(0.0109)	(0.0105)	(0.0157)	(0.0151)
FRA greater than age 65	0.0099	-0.0013	0.0031	-0.0258
	(0.0126)	(0.0127)	(0.0182)	(0.0183)
Husband's FRA greater than age 65	0.0250*	0.0181	0.0493**	0.0255
	(0.0149)	(0.0158)	(0.0215)	(0.0227)
Combined effect of husband and wife less than age 70 in 2000	-0.0094	-0.0065	-0.0029	0.0005
	(0.0120)	(0.0116)	(0.0174)	(0.0167)
Number of observations	1,625	1,522	1,625	1,522
B. Women first observed age 75 or older				
Less than age 70 in 2000	0.0000	0.0029	0.0288	0.0382**
	(0.0113)	(0.0116)	(0.0180)	(0.0183)
Husband less than age 70 in 2000	0.0138	0.0091	0.0058	-0.0021
	(0.0116)	(0.0118)	(0.0186)	(0.0187)
Husband's FRA greater than age 65	0.0407	0.0432	0.1242**	0.1327***
	(0.0322)	(0.0324)	(0.0514)	(0.0511)
Combined effect of husband and wife less than age 70 in 2000	0.0138	0.0120	0.0346*	0.0361*
	(0.0115)	(0.0118)	(0.0183)	(0.0186)
Number of observations	1,050	1,003	1,050	1,003

Notes: See notes to Tables 1, 2, 3, 6, and 9. Poverty is based on the sum of adjusted Social Security benefits and other income relative to the poverty line (all in 2013 dollars). Note that the cell sizes that identify the coefficient of "Husband's FRA greater than age 65" are very small, given that the women have to be 75 or older, and the men cannot turn 65 until 2003.For the age 75 sample, there are 26 women with husband's whose FRA is greater than age 65. Five of these women live below 200 percent of the poverty line.

Table 12: The Effect of the 2000 Elimination of the Retirement Earnings Test on Poverty, Women with Husband Observed, Women First Observed Age 70 or Older and 75 or Older Samples, Two-Stage Least Squares Estimates

	Born 1918-1942	Born 1925-1940	Born 1918-1942	Born 1925-1940
Dependent variable:	Living below 150% of poverty line	Living below 150% of poverty line	Living below 200% of poverty line	Living below 200% of poverty line
	(1)	(2)	(3)	(4)
A. Women first observed age 70 or older				
Age at claiming benefits in months (endogenous)	0.0009	0.0002	0.0020	0.0017
	(0.0011)	(0.0005)	(0.0016)	(0.0016)
Husband's age at claiming benefits in months (endogenous)	0.0001	-0.0000	-0.0018	-0.0019
	(0.0013)	(0.0006)	(0.0019)	(0.0018)
FRA greater than age 65	0.0081	0.0031	0.0038	-0.0269
	(0.0121)	(0.0059)	(0.0176)	(0.0178)
Husband's FRA greater than age 65	0.0263*	-0.0030	0.0464**	0.0217
	(0.0158)	(0.0079)	(0.0230)	(0.0237)
Combined effect of husband and wife less than age 70 in 2000	0.0010	0.0002	0.0002	-0.0002
	(0.0013)	(0.0007)	(0.0019)	(0.0020)
Instruments	Less than age 70 in 2000, husband less than age 70 in 2000			
Minimum eigenvalue statistic	19.86 (7.03)	18.48 (7.03)	19.86 (7.03)	18.48 (7.03)
Number of observations	1,625	1,522	1,625	1,522
B. Women first observed age 75 or older				
Age at claiming benefits in months (endogenous)	0.0001	-0.0003	-0.0027	-0.0043**
	(0.0011)	(0.0013)	(0.0018)	(0.0022)
Husband's age at claiming benefits in months (endogenous)	-0.0016	-0.0011	-0.0011	-0.0004
	(0.0013)	(0.0013)	(0.0020)	(0.0022)
Husband's FRA greater than age 65	0.0441	0.0468	0.1378***	0.1511***
	(0.0323)	(0.0324)	(0.0524)	(0.0535)
Combined effect of husband and wife less than age 70 in 2000	-0.0015	-0.0014	-0.0038*	-0.0048*
	(0.0013)	(0.0015)	(0.0020)	(0.0024)
Instruments	Less than age 70 in 2000, husband less than age 70 in 2000			
Minimum eigenvalue statistic	15.93 (7.03)	14.35 (7.03)	15.93 (7.03)	14.35 (7.03)
Number of observations	1,050	1,003	1,050	1,003

Notes: See notes to Tables 1, 2, 3, 6, and 11. First-stage estimates are not reported.

Appendix Table A1: Percentage of Women Age 62 or Older by Type of Entitlement and Year

	Entitled worker only	Dually-entitled		Spouse only	Survivor only
		Spouse	Survivor		
1995	36.2	11.5	14.4	14.3	23.6
2000	38.0	12.0	15.6	12.9	21.5
2005	41.4	12.0	16.0	11.4	19.3
2010	46.3	12.1	15.5	9.6	16.4
2011	47.5	12.0	15.3	9.3	15.9
2013	48.7	11.9	15.1	9.0	15.3

Source: Data from the Social Security Administration's *Annual Statistical Supplement, 2013*, Table 5.A14.

Appendix Table A2: Discrepancies between Claiming Ages/Dates in Self-Reported and Linked Administrative Data

Diff = Public – Admin.	Full sample (1)	Respondents born before 1931 (2)	Respondents born in 1931 or later (3)
Less than -12 months	30 or fewer	16	15 or fewer
-4 to -12 months	85	22	63
-1 to -3 months	237	60	177
No difference	812	221	591
1 to 3 months	1134	174	960
4 to 12 months	478	221	257
13 or more months	299	190	109
Mean	4.1	8.8	2.1
Standard Deviation	11.1	16.8	6.7

Notes: The sample in the table contains men who have agreed to have their Social Security records linked to their HRS survey responses. The sample is limited to those individuals who have never claimed disability benefits and whose current and initial claiming type is equal to a retired worker benefit. Any men who are receiving a second benefit based on another individual's account, whose beneficiary date of birth is not equal to the account holder's date of birth, whose "historical current entitlement code" is equal to disabled, whose date of current entitlement does not equal the date of initial entitlement, or whose age at claiming in the administrative data is calculated to be less than age 62, are removed from the sample. Finally, the sample is limited to those whose self-reported age at claiming (in months) is between 744 and 851 (inclusive).

Appendix Table A3: The Effect of the 2000 Elimination of the Retirement Earnings Test on Benefit Entitlement, Song and Manchester (2007) Replication, Probit and OLS Estimates

Song and Manchester (2007) results:	Between 0.0219 and 0.0502			
	OLS		Probit	
	Public data (1)	SSA admin. data (2)	Public data (3)	SSA admin. data (4)
Attained ages 65-69	0.0469***	-0.0082	0.0669***	0.0021
	(0.0102)	(0.0106)	(0.0113)	(0.0093)
Number of observations	19,644	13,779	19,644	13,779

Notes: The samples include both men and women. The Song and Manchester results presented in the table are from Table 4, p. 686, Model II (the base model). In their analysis, Song and Manchester (2007) use two separate control groups. One group of the treatment groups includes those who had attained ages 65 to 69 by January 1 of the calendar year. The second treatment group includes those who turn age 65 during the year. The younger control group contains those who turn age 62 to age 64 during the year. The older control group contains those who have attained ages 70 to 72 as of January 1 of the calendar year. Their regressions include interaction of treatment status with year dummies, dummies for being in the younger or older control group, calendar year dummies, an indicator for whether the person-year observation is male, and an indicator for whether the person-year observation is white. For probit estimates, marginal effects evaluated at sample means and varying the treatment variable from zero to one are reported.