Labor Supply Effects of the Recent Social Security Benefit Cuts: Empirical Estimates Using Cohort Discontinuities

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June 2005

Abstract

In response to an earlier "crisis" in Social Security financing two decades ago the Congress implemented an increase in the Normal Retirement Age (NRA) of two months per year for cohorts born in 1938 and afterward. These cohorts reached retirement age in 2000 and in this paper I study the effects of these benefit cuts on recent retirement behavior. The evidence strongly suggests that the mean retirement age of the affected cohorts has increased by around half as much as the increase in the NRA. If these increases in work effort by older workers continue, it will have extremely important implications for the estimates of Social Security trust fund exhaustion that have played such a major role in recent discussions of Social Security reform.

I am particularly indebt to Orley Ashenfelter. I would also like to thank Wioletta Dziuda, Bo Honoré, Alan Krueger, Franco Peracchi and all participants at the Princeton Labor Lunch Seminar.

Keywords: normal retirement age, retirement behavior, social security reform

JEL classification codes: H55, J26

1 Introduction

In 1983 the US Congress implemented an increase in the Normal Retirement Age (NRA) of two months per year for cohorts born in 1938 and afterward. Due to the way Social Security benefits are computed, each 2-months increase in the NRA translates into a little bit more than a 1 percentage point reduction in Social Security benefits. The cohorts affected by these benefit cuts reached the early retirement age (62) in 2000 and in this paper I study the effects of these benefit cuts on recent retirement behavior.

The normal retirement age is gradually rising from 65 to 67. Workers born in 1938 face a normal retirement age of 65 years and 2 months, which means that, everything else equal, their benefits are 1 percent smaller than those paid to workers born before 1938, with a NRA of 65. The 1939 cohort has a NRA of 65 and 4 months and so on up to the 1943 cohort, whose NRA will be 66. After a 10-year break, the reform will resume with cohort 1955, reaching an NRA of 67 for workers born after 1960.

Increasing the NRA is likely to influence two important decisions that a worker faces at the end of her career: (a) when to start collecting Social Security benefits, and (b) when to retire. Since benefits are adjusted actuarially fairly with respect to the entitlement age, the long-term solvency of the Social Security trust fund can potentially only be improved by the decision about retirement. An increase in labor force participation generates more contributions, which are the main source of revenue for the trust fund.

In this paper I provide the first ex-post evaluation of what the actual effects of increases in the NRA have been. The evaluation may provide both substantive evidence to guide further reforms and a guide to calibration of structural models of retirement decisions. My results suggest that previous research may have seriously underestimated the influence of a change in the NRA on retirement behavior and raise a hustings question about how best to improve these models. Moreover, the advantage of using the change in the NRA to estimate the effect of Social Security incentives on labor supply is that the exact change in benefits is known, not prone to measurement error, and is exogenous.

Due to the timing of the reform I treat the 1937 birth cohort as the control group and the post 1938 cohorts, which face a reduction in benefits, as the treatment groups in the analysis. Figure 1 plots the fraction of, respectively, male and female workers in the sample who are retired form the labor force as a function of their age, both before and after the changes in the NRA. I argue that the most obvious candidate for the distance between the two CDFs of the retirement age is the increase in the NRA.

In terms of changes in the average retirement age the point estimates generally imply, depending on the model used, an increase in the actual age of retirement of around 35 percent to 75 percent of the increase in the NRA.

My estimates are more than three times as large as previous out-of-sample predictions. These predictions suggested that the labor supply response to the change in the NRA would be small, though huge potential uncertainty exists about such predictions. Coile and Gruber (2000) simulate the effects on retirement of a one year increase in the NRA. The effect in the 61 to 65 age range is estimated to reduce the age of retirement by between one-half and two months. Another recent paper that contains simulations of a one year increase in the NRA is a report for the SSA written by Panis, Hurd, Loughran, Zissimopoulos, Haider and St.Clair (2002). Their estimate on the average retirement age is around seven days. Both studies had to rely on estimates based on the cross-sectional variation in labor supply that could be related to differences in Social Security benefits. One problem is that Social Security benefits depend on the whole history of wages, and so endogeneity can be an issue. Moreover, present discounted values of future streams of benefits are likely to be measured with error, which may downward bias their estimates.

Despite the reforms passed in 1983, the trust fund is projected to become insolvent in less than forty years. While this date of insolvency is often portrayed by the news media as certain, a great deal of uncertainty surrounds these estimates. One of the most important sources of uncertainty is the behavior of future workers and retirees.[1] To make better predictions it is important to understand how workers' behavior may have been affected by the increase in the NRA.

Section 2 introduces a simple intertemporal model of retirement. It's main purpose is to highlight that in theory transitional effects that arise when benefit cuts are unexpected can generate big changes in labor supply. Section 3 presents the model used for the empirical estimation. Results are shown in section 4, while section 5 concludes. Appendix A describes the data I use.

2 A simple intertemporal model of retirement

The life-cycle theory predicts that a worker's reaction to benefit cuts, which is a decrease in lifetime income, will depend on when she first learns about the reform. Attentive workers may have started reacting to the reform in 1983, and for these workers the change in retirement behavior is likely to be small due to 20 years of consumption smoothing. Some workers may have learned about the increase of the NRA in 1995, when the SSA started mailing a Social Security Statement to all workers age 60 and over. The statement shows estimated benefits at different ages of retirement, including the first possible age of retirement and the NRA. In 2000 the SSA also added a special insert to the statement containing the changes in the NRA. Very distracted workers may learn about the benefit at the time they claim the benefits.

The purpose of the model is to show what the expected reaction in terms of both consumption and retirement is as a function of the date at which the worker is informed about the benefit cut. This model is quite standard.[2] It assumes that workers maximize their utility over consumption (C) and the time of retirement (z). Retirement is an

[1]See Anderson, Lee and Tuljapurkar 2003

[2]See for example Colombino 2003 or the working paper version of Hurd, Smith and Zissimopoulos 2002

absorbing state and workers claim benefits at the time they retire and face a perfect capital market rate of return r. There is no uncertainty about wages W and mortality. The worker's problem takes the following form:

$$\max_{z,C_t} V(z) = \int_0^z e^{-\delta t} U_W(C_t)dt + \int_z^D e^{-\delta t} U_R(C_t)dt \qquad (1)$$

s.t.

$$\int_0^D e^{-rt} C_t dt = \int_0^z e^{-rt} W_t dt + \int_z^D e^{-rt} R_t dt \ , \qquad (2)$$

where D is the date of death. In order to obtain closed form solutions, I assume that the utility function is $U_R(C) = \ln C$. Disutility from work is captured by an additive constant $U_W = U_R - \epsilon$, where U_W is a worker's utility level and U_R is that worker's utility in retirement. In this setup, e^{ϵ} represents the factor by which the worker's consumption has to be increased to generate the same level of a retirees' utility. Notice that what is called disutility from work may additionally capture the observation that retirees tend to make better consumption choices (Aguiar and Hurst 2004) and that retirees do not have work related costs. I further assume for simplicity that the rate of preference equals the interest rate, $\delta = r$ and real wages are constant over time, $W_t = W$. The benefit formula used by the SSA expresses benefits as a function of past wages and increases with the age of retirement, z:

$$R(z, W) \quad = \quad R(W)(1 + g(z - NRA)) \ ,$$

where g represents the actuarial adjustment factor.

In Appendix B I show that this simple model gives two important predictions. First, for reasonable parameters, increasing the NRA delays retirement and reduces consumption (see appendix). This result implicitly assumes that the Social Security rules change at time zero, when the worker starts working. Second, for reasonable parameters, if the rules change when the worker is already working, the reaction in terms of consumption and retirement is stronger. This occurs because an early-informed worker has more time to smooth consumption over time, and thus will not postpone retirement as much as a late-informed one.

3 Empirical Strategy

The estimation strategy is to measure the distance between the cumulative distribution functions (CDFs) of the retirement age of workers with different NRAs.

I estimate the following linear model by least squares:

$$y_i = \sum_{a=61}^{65} 1(A_i = a)\left(\alpha_a + \sum_{c \neq 1937} \beta_{a,c} 1(C_i^* = c)\right) + \gamma' X_i + \epsilon_i \ , \qquad (3)$$

where y_i is equal to 1 when the worker is retired and zero otherwise. $1(A_i = a)$ is equal to 1 if the worker is a years old and 0 otherwise and $1(C_i^* = c)$ is equal to 1 if the worker is born in year c and 0 otherwise.

Since I include all age dummies and I omit the 1937 cohort dummy and the constant term, the $\beta_{a,c}$ measures the difference between cohort c and cohort 1937 at age a in the CDFs of the retirement age, $\widehat{\beta}_{a,c} = E[Y|C = c, a, X = 0] - E[Y|C = 1937, a, X = 0]$.

One problem with evaluating this reform is that using CPS data the year of birth variable can sometimes be misclassified. CPS data contain a precise measure of the age of the respondent in the survey week,[3] but this information coupled with the information of the

[3]CPS respondents provide their date of birth, though this information is later discarded from the public-use data. Unfortunately, because of the weak follow-up and the noisy identification of observations across waves, using the longitudinal component of the CPS allows me to get an exact measure of the year of birth for only a small number of observations. To match observations over time I use the conservative

survey year provides at least an imperfect measure of the year of birth. Misclassification errors are not uncommon in empirical research. Gruber and Orszag (2003) in a paper that analyzes the impact of the earnings test on labor supply take the most conservative approach of deleting observations for which ambiguity exists about the earnings test regime. Krueger and Pischke (1992) warn the reader about the a probability of misclassification of around 20 percent when using the March CPS to establish the year of birth, but they do not correct for that.

Since months of birth are approximately uniformly distributed (Table 3), the probability of misclassifying the year of birth based on the survey month is known. If I take the simple approach of generating the birth cohort as the difference between the survey year and age, $cohort = year - age$, for example in the January survey the probability of misclassifying someone who was born in 1936 as a 1937 cohort is (assuming that the survey is carried out in the last day of the month) around 11/12. The reason is that someone surveyed in January is very likely to have his birthday later in the year. The probability of misclassification is 10/12 in February, and, carrying out the calculation, zero in December. The probability of misclassification using this method would be on average equal to one half.

A better way to assign the birth year is to minimize the probability of misclassification. Adding a year to the cohort if the survey month is one of the first six months of the year causes the average probability of misclassification to drop to one quarter. When I use this definition I call it the naive method. When I use this method and I additionally restrict the sample to the January and December surveys the probability of misclassification is only one over twelve.[4] I call this method the restricted method.

There is an obvious trade-off between minimizing the probability of misclassification

approach of first matching by the CPS identifiers (hrhhid huhhnum hurespl), race and gender. Whenever, after this first step the standard deviation of age is bigger than one half, I additionally match by education, which for elderly people is normally constant over time (Madrian and Lefgren 1999).

[4]To be more precise, given that the survey week always contains the 19th of the month, the probability is 19/365 in January and 11/365 in December.

and maximizing the statistical power. In order to work with the whole sample, instead of minimizing the error, I can use the information about the probability distribution of the misclassification errors and estimate a model with misclassification (Aigner 1973). The only empirical paper I am aware of that uses a similar approach is Card and Krueger (1992). Let $Y \in \{0, 1\}$ be 1 if the worker is retired and define C^* to be the true cohort and C the observed cohort. The misclassification probabilities are known and assumed to depend only on the survey month m, $p(m) = \Pr(C^* = c - 1|C = c, m)$. If education or other regressors are correlated with the month of birth,[5] the estimator may not be consistent. $\Pr(Y = 1|C = c, m, a, X) = E[Y|C = c, m, a, X]$ represents the conditional probability of being retired by age a, given that a worker is observed in month m to be born in year c, while $E[Y|C^* = c, m, a, X]$ represents the probability of being retired given that a worker is truly born in year c. For ease of notation I will discard the other independent variables X, but probabilities that are not misclassification probabilities are supposed to be conditional on X.

Assuming that given the true cohort the mismeasured one is not informative, I have that

$$E[Y|C = c, C^* = c, m, a] = E[Y|C^* = c, m, a] \ .$$

By the law of total probability

$$E[Y|C = c, m, a] = (1 - p(m))E[Y|C^* = c, m, a] + p(m)E[Y|C^* = c - 1, m, a] \ .$$

The probability of being retired depends on the survey month as well, since, conditional

[5] See Angrist and Krueger (1991).

on a birth year (the true or the observed one), later in the year workers tend to be older. Assuming that conditional on the cohort C^*, the dependency on the survey month is additively separable and does not change across cohorts, $E[Y|C^* = c, m, a] = E[Y|C^* = c, a] + g(m, a)$. If all workers were retiring only in January $g(m, a)$ would be 0. Plugging this into the previous equation, I get that

$$E[Y|C = c, m, a] = (1 - p(m))E[Y|C^* = c, a] + p(m)E[Y|C^* = c - 1, a] + g(m, a)$$

Averaging over the different survey months after defining $p = \sum_m p(m) \Pr(M = m)$ results in

$$E[Y|C = c, a] = (1 - p)E[Y|C^* = c, a] + pE[Y|C^* = c - 1, a] + g(a) \ ,$$

where $g(a) = E(g(m))$ The main reason for specifying the dependence of retirement on the survey is to remember that in the empirical analysis it is important to keep a similar distribution of survey months when comparing different cohorts. Having this in mind, if all months of the year are included in the empirical analysis, from the definition $E[Y|C^* = c, m, a] = E[Y|C^* = c, a] + g(m, a)$ it follows that $g(a)$ is zero.

Solving for the true effect I get a recursive formula, where the true probability of being retired for a given cohort is a function of the observed probability and the true probability of the previous cohort,

$$E[Y|C^* = c, a] = \frac{E[Y|C = c, a] - E[Y|C^* = c - 1, a]p}{1 - p} - g \ . \tag{4}$$

At this point I need a starting point for the recursion. I use $E[Y|C = 1935, a] =$

$E[Y|C^* = 1936, a]$ as a staring point, which implies that $E[Y|C^* = 1936, a] = E[Y|C = 1936, a]$. This allows me to analyze the differences in the CDF between 2 pre-reform cohorts, the 1936 and the 1937 one.

It can be shown that this recursion can be implemented by estimating the following linear model

$$y_i = \sum_{a=61}^{65} 1(A_i = a) \left(\sum_{c=1936}^{1939} \gamma_{a,c} \Pr(C_i^* = c) \right) + \gamma' X_i + \epsilon_i \ , \tag{5}$$

with the initial condition $\Pr(C_i^* = 1936) = \{0, 1\}$.

Conditioning on c, a, and $X = 0$:

$$\begin{aligned} E[Y|C = c, a, X = 0] &= \gamma_{a,c} \Pr(C^* = c|C = c) + \gamma_{a,c-1} \Pr(C^* = c - 1|C = c) \\ &= \gamma_{a,c}(1 - p) + \gamma_{a,c-1} p \end{aligned} \tag{6}$$

Rearranging terms,

$$\gamma_{a,c} = \frac{E[Y|C = c, a, X = 0] - \gamma_{a,c-1} p}{1 - p} \ , \tag{7}$$

which resembles Eq. 4.

Instead of estimating $\widehat{\gamma}_{a,c}$ I estimate $\widehat{\beta}_{a,c} = \widehat{\gamma}_{a,c} - \widehat{\gamma}_{1937,c}$ using eq. 3 with the only difference that $\Pr(C_i^* = c)$ substitutes for the previous $1(C_i^* = c)$.

The $\beta_{a,c}$s measure the difference between the cohorts' cumulative distribution functions that I showed in the figures, but also control for differences that may be due to observable characteristics X. It is assumed that adjacent birth cohorts do not differ by unobservable factors that affect retirement behavior.

A more easily interpretable result can be obtained from the sum the estimated coefficients. If outside of the 62–65 window the CDFs of the different cohorts are either assumed to be identical or unrelated to the change in the NRA, this sum measures the effect of the reform on the average retirement age. For the 1938 cohort, for example,

$$
\begin{aligned}
E[A_{38}] - E[A_{37}] &= \sum_{a=62}^{66} a[\Pr_{38}(A=a) - \Pr_{37}(A=a)] \\
&= \sum_{a=62}^{66} a[(\Pr_{38}(A \leq a) - \Pr_{37}(A \leq a)) - (\Pr_{38}(A \leq a-1) - \Pr_{37}(A \leq a-1))] \\
&= \sum_{a=62}^{66} a(\beta_{a,38} - \beta_{a-1,38}) \\
&= 62(\beta_{62,38} - \beta_{61,38}) + ... + 66(\beta_{66,38} - \beta_{65,38}) \\
&= 62(\beta_{62,38} - 0) + ... + 66(0 - \beta_{65,38}) \\
&= -\sum_{a=62}^{65} \beta_{a,38} \; . \qquad (8)
\end{aligned}
$$

Notice that for men the differences outside of the 62–65 window are indeed smaller (Figure 2), while for women the difference at age 61 are quite large (Figure 3). By not including differences outside the 62–65 window I may underestimate the effect on the average retirement age.

Tables 1 and 2 contain the summary statistics of the two samples that are later used in the analysis. On average the observed birth cohorts are quite similar.

4 Estimation Results

First, I divide workers according to whether they were born before or after 1938 based on the naive method. Figure 1 shows the corresponding CDFs for the female (right panel) and the male sample (left panel). The primary alternative hypothesis is that this gap is a product of a preexisting trend towards later retirement. As Quinn (1999) showed, the trend towards early retirement stopped in the late 80s and early 90s. However, if there was a trend towards later retirement that is unrelated to the benefit cuts, there would be a significant difference in labor force participation rates across all birth cohort, while Figures 2 and 3 show that the CDFs of workers born between 1935 and 1937 are very similar. Instead, the increase in labor force participation starts with workers born in 1938 or later, and these are precisely the workers who experienced benefit cuts.

Next, I measure the difference between the CDF's by estimating equation (3). I estimate the model separately for men and women, but the estimates related to the female sample have to be interpreted with some additional caution. For the majority of women the Social Security benefits based on their own earnings history are smaller than 50 percent of the benefits they could get by applying for spouse's benefits. It is unclear how much these women react to changes in their spouse's benefits.

The estimated distance between the cumulative distribution functions ($\widehat{\beta}_2$) are shown in Tables 4 and 5. For each of the three models columns (1), (3) and (5) contain only age and cohort dummies, while in the columns (2), (4) and (6), additionally control for marital status, education, race, children in the household, total members of the household, geographic region, veteran status and whether the household resides in a metropolitan area. Controlling for these variables generally slightly reduces the estimated effects. Table 6 shows the sum of the estimated coefficients, the sample equivalent of equation (8). These estimates represent the change of the average retirement age with respect to the cohort of 1937.

The first striking result from Tables 4 and 5, is that for all three models and for both, men and women, the estimated difference in CDFs between, on one side, the 1938 and 1939 cohorts, and on the other side, the 1937 cohort are almost always negative, meaning that in the age range 61–65 the CDF of the 1937 cohort lies above the CDF of the other two cohorts. Although not all post–reform $\widehat{\beta}$s are significant, their sums are, apart from few exceptions, highly significant (Table 6), which suggests that the increase in the NRA generated an increase in the reform had a significant effect on the average retirement age. On the other hand the 1936 and 1937 CDF are quite similar, and this translates into changes in the average retirement age that is not significantly different from 0.

Due to the misclassification error, the estimates of the naive model are generally smaller, while the sophisticated method produces effects that tend to lie in between the other two. The true effect is likely to be underestimated by the naive method, reason why I will restrict the analysis to the remaining two methods.

Women born in 1938, who face an increase of 2 months of their NRA, have a change in their average retirement age that lies between 1 and 1.4 months (8.16 percent and 11.18 percent of a year). The female 1939 cohort, whose NRA increased by 4 months, shows a change that lies between 1.1 and 2.3 months. Men born in 1938 show an effect of 1.6 to 1.9 months, while the one born in 1939 show an effect that lies between 1.4 and 1.6 months. From this one might conclude that the effect does not seem to increase for later cohorts, at least for men. Looking back at figures 2 to 5, though, seems to suggest otherwise. The 1940 cohort, not included in the regressions because it is to early to observe the cohort at 65, has lower retirement rates between ages 62 and 64.

4.1 Alternative explanations

The identification is based on the assumption that the observed change in labor force participation across contiguous birth cohorts is due to the change in the NRA. The fact that for the 1938 and 1939 cohorts the estimated β_2s are negative at all ages allows me to rule out that single shocks are driving all the results. Take for example the stock market crisis of 2001. Workers with defined contribution plans may react to such shocks by working longer to make up for the financial losses. While this may explain single differences in the cumulative distribution functions, in particular the 1939–1937 one at age 62 or the 1938–1937 at age 63, it cannot explain the other differences, where cohorts are affected in a similar way. Notice also that at the time of the 2002–2003 stock market crisis the youngest cohort (1940) is already 63 years old. Unless the effect related to the stock market crisis is heterogenous across age, it will difference out when summing up the coefficients to get the effect on the average retirement age. Moreover, Coile and Levine (2004) find "no evidence that changes in the stock market drive aggregate trends in labor supply." This is mainly due to fact that, although 45 percent of all workers are covered by a pension plan, few of them have substantial stock holdings.

Another possible confounding effect is the 2000 Earnings Test removal above the NRA. Earnings of Social Security beneficiaries above the earnings test threshold, up to their benefit amount, are taxed away at a 50 percent rate between age 62 and 65, and, before 2000, at a 33 percent rate between 65 and 69. While the benefit that are taxed away due to the earnings test are not lost, but rather postponed at an actuarially fair rate, there is some evidence that people perceive the earnings test as a pure tax.[6] Since my analysis is focused on ages 61 to 65, this change may affect my results only if there are spillover effects on younger ages, in other words only if workers decide to continue working in order to reach the age at which they can work without being taxed. Analogous to the previous case, this would only affect single differences, particularly the 1939-1937 one at age 61 or the 1938-1937 at age 62, and only if spillover effects reach 4-5 years back!

In order to exclude that results are driven by labor market shocks, the same equation has been estimated using weekly hours of work as the dependent variable. There are no significant differences in hours of work across these cohorts. The results are also not driven by differences in part-time work or disability status. Including disabled workers, or part-time workers (work less than 35 hours) to the retired workers also does not alter

[6]See Gruber and Orszag 2003.

the results (results available upon request).

5 Conclusions

An aging population and low labor force participation rates for older workers have worsen the financial situation of the Social Security Trust Fund. Aware of this, some twenty years ago several reforms were passed on the recommendation of the Greenspan commission. Their aim was to cut benefits and increase labor force participation. Among other changes, the reform scheduled an increase in the normal retirement age for workers born after 1938.

I find evidence that workers reacted very strongly to this increase in the NRA. The average retirement age for the 1938 and 1939 birth cohorts increased by around 1.5 months. Given that the change in the NRA was for these two cohorts, respectively, 2 and 4 months, this represents a large effect. One puzzling observation is, however, that the 1939 cohort reacted in the similar way as the 1938 cohort. On the other hand, looking at the CDF of the 1940 cohort, not included in the regressions because I do not observe that cohort at age 65, it seems that that cohort exhibits larger changes in the average retirement age. To calculate a precise estimate the analysis of this paper has to be repeated in a few years. Given that there is an ongoing intense work on reforming Social Security, conducting early analysis with limited data is, I believe, essential.

Previous studies, using out-of-sample predictions, have estimated much smaller effects on labor force participation. Beside the fact that these estimates are based on older cohorts that may respond differently to retirement incentives, 3 major factors may have downward-biased previous results. The first is that these models are not capturing any effect due to norms that may be related to the NRA. There is strong evidence that workers may look at the NRA as a focal point. The second is that estimates based on these models, since benefits are a function of past earnings, may suffer from endogeneity bias. The third is that these models, since they are estimated using cross-sectional variation in Social Security benefits and retirement status, may capture long-term effects, while the 1983 reforms may have been unexpected. Using a simple intertemporal model of retirement I show how this can generate larger changes in the average retirement age than would otherwise be expected. The third problem is that in order to construct Social Security wealth, a component of all forward-looking incentives to retire, the researcher needs detailed information about past and future earnings, interest rates, and preferences; in short, measurement error may be an issue. The increase in the NRA generates a clear reduction in Social Security Wealth that is free of measurement error.

Despite the 1983 reform, the Social Security trust fund is projected to become insolvent in around 40 years. The Social Security projections are only one of several projections made by other institutions. A common feature of all this projections is that they depend heavily on the way the future behavior is modelled. My results may help to evaluate the importance of the increase of the NRA on labor force participation.

According to the 2003 Technical Panel on Assumptions and Methods (*Technical Panel on Assumptions and Methods* 2003) little documentation is available on the trustees' way of forecasting labor force participation. The same panel explains that the base methodology is based on three steps: The first is to estimate autoregressive "age, sex, marital status, and presence of children" specific labor force participation rates models that control for economic, demographic, and policy variables. For older people instead of LFPRs hazard rates are used, and Social Security benefits (relative to past earnings) and the fraction of workers affected by the Social Security earnings test are included in the regressions, though it is not clear how big the age groups are. The second step is to subjectively adjust some estimated coefficients based on economic theory, prior beliefs, and the "full mosaic" of all estimated models. The last step is to estimate fitted values based on projections of explanatory variables. This model is likely to be accurate in case of smooth changes over time. The problem is that the increase in the NRA may have introduced a break in the trend at the very end of the time span used by the trustees, and therefore difficult to detect, especially if age groups are merged together. According to the 2004 Trustees report (*The 2004 Annual Report* 2004)

> "... changes in available benefit levels from Social Security and increases in the normal retirement age, and the effects of modifying the earnings test are expected to encourage work at higher ages. Some of these factors are modelled directly."

The 2000 Technical Panel on Assumptions and Methods (*Technical Panel on Assumptions and Methods 2003)* recommends that "Social Security should be considered explicitly since it may result in higher participation rates." If the increase in NRA affects workers born after 1939 the way it seems to have affected the 1938 and 1939 birth cohorts, the trustees should definitely follow this recommendation.

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A Data

I use the CPS monthly data from January 1998 to December 2004. The CPS data contain information about the respondent's age by the end of the survey week, usually the second week of the month.[7]

I restrict the data to individuals born between 1936 and 1939, aged 61 to 65. While the upper limit 65 is chosen because it is the last available age for the 1939 cohort, the lower limit of 61 is chosen to avoid attributing differences before that age to the change in the NRA. Workers who retire early need to wait at least until age 62 in order to claim the benefits. Differences in retirement rates before 62 are therefore unlikely to be related to the increase in the NRA. Including age 61 allows for possible small spillover effects. These restrictions represent conservative choices and may understate the overall effect, though, since, as will be shown later, differences in retirement rates under age 62 and above age 65 tend to be smaller, the bias is likely to be small. I also discard data on individuals who are not in the labor force and do not report to be retired. This group includes people with disabilities. The CPS has a much bigger sample size than the Health and Retirement Survey (HRS). For each 1937-1939 birth cohort, aged between 61 and 65 there are around 60.000 observations, while the HRS contains only 1000 observations for people born in 1937 and aged 61 to 63. Also, the last available HRS has been collected in 2002, when workers born in 1938 were only 64.

The disadvantage of these data is that there is no information on Social Security insured status and on pension benefits. Fortunately almost all active and retired men and women above 62 are eligible for Social Security benefits (Panis, Hurd, Loughran, Zissimopoulos, Haider and St.Clair 2002), so that the reform is likely to affect almost the entire cohort of workers born after 1938. This allows me to estimate labor supply responses to benefit cuts without actually observing individual benefits.

[7]The reference week for CPS is the week (Sunday through Saturday) of the month containing the 12th day.

All the analysis uses unweighted data. When I use CPS weights, the distance between the distribution functions is, only for men, slightly smaller. The reason for this may lie in the revision to the weighting procedures, and the switch from the 1990 to the 2000 Census, that, according to the Bureau of Labor Statistics, affected the comparability of CPS data series over time (Bowler, Ilg, Miller, Robison and Polivka 2003). Since this changes fall exactly in the middle of the sample, the analysis is carried on without using CPS weights.

B The intertemporal model or retirement

The first order conditions of the model are:

$$\begin{aligned} dz &: U_W(C_t) = U_R(C_t) - \mu(W_z - R_z(z) + \int_z^D e^{r(z-t)} \frac{\partial R_t(z)}{\partial z} dt) \\ dC &: \frac{\partial U_x(C_t)}{\partial C_t} = \mu \quad x = W, R \end{aligned}$$

Given these assumptions, the system of equations that define the equilibrium is:

$$\epsilon C = W - R(1 + \frac{.05}{10}(z - NRA)) + R\frac{.05}{10}(\frac{1}{r} - \frac{1}{r}e^{r(z-D)})$$

$$\begin{aligned} C &= \frac{1 - e^{-rz}}{1 - e^{-rD}} W + \frac{e^{-rz} - e^{-rD}}{1 - e^{-rD}} R(1 + \frac{.05}{10}(z - NRA)) \\ &= \alpha(z)W + (1 - \alpha(z))R(1 + \frac{.05}{10}(z - NRA)) \end{aligned}$$

Totally differentiating:

$$\begin{pmatrix} 1 & \frac{re^{-rz}}{1-e^{-rD}}((1+\frac{.05}{10}(z-NRA))R-W)-\frac{.05}{10}R\frac{e^{-rz}-e^{-rD}}{1-e^{-rD}} \\ \epsilon & \frac{.05}{10}R(1+e^{r(z-D)}) \end{pmatrix} \begin{pmatrix} dC \\ dz \end{pmatrix}$$
$$= \begin{pmatrix} -\frac{.05}{10}R\frac{e^{-rz}-e^{-rD}}{1-e^{-rD}} \\ \frac{.05}{10}R \end{pmatrix} dNRA$$

and solving:

$$\begin{pmatrix} \frac{dC}{dNRA} \\ \frac{dz}{dNRA} \end{pmatrix}$$
$$= \frac{1}{\Delta} \begin{pmatrix} .00\,5R\left(-1+e^{-rD}\right)\left(1+e^{-r(-z+D)}\right) & re^{-rz}(R-W)+.00\,5Re^{-rz}(rz-rNRA+e^{r(z-D)}-1) \\ -\epsilon\left(-1+e^{-rD}\right) & \left(-1+e^{-rD}\right) \end{pmatrix}$$
$$\begin{pmatrix} -\frac{.05}{10}R\frac{e^{-rz}-e^{-rD}}{1-e^{-rD}} \\ \frac{.05}{10}R \end{pmatrix}$$

where

$$\Delta = \frac{.05}{10}R(\left(1+e^{-r(-z+D)}\right)\left(-1+e^{-rD}\right)+\epsilon e^{-rz}(r(z-NRA)+e^{r(z-D)}-1)$$
$$-\epsilon re^{-rz}(W-R)\ .$$

Notice that if $r(z-NRA)+e^{r(z-D)}-1<0$, then $\Delta<0$. The first expression can only be positive if at the same time the worker retires after her NRA ($z>NRA$) and the interest rate is extremely large. It follows that for reasonable parameters the retirement age increases when the NRA increases,

$$\frac{dz}{dNRA} = \frac{\frac{.05}{10}R}{\Delta}(-\epsilon\left(e^{-rz}-e^{-rD}\right)-1+e^{-rD})>0\ , \qquad (9)$$

while consumption decreases if,

$$\frac{dC}{dNRA} = \frac{\left(\frac{.05}{10}R\right)^2}{\Delta} e^{-rz}\left(e^{r(z-D)}\left(1-e^{r(z-D)}\right)+r\left(\frac{R-W}{\frac{.05}{10}R}+z-NRA\right)\right) < 0.$$

or

$$e^{r(z-D)}\left(1-e^{r(z-D)}\right)+r\left(\frac{R-W}{\frac{.05}{10}R}+z-NRA\right) > 0\ .$$

Notice that the first term is always positive, while the second is not. Now assume that an increase of NRA to NRA' has not been anticipated. Up to time z the worker behaves like in the previous case

$$\epsilon C = W - R(1+\frac{.05}{10}(z-NRA)) + R\frac{.05}{10}(\frac{1}{r}-\frac{1}{r}e^{r(z-D)})$$

$$C = \frac{1-e^{-rz}}{1-e^{-rD}}W + \frac{e^{-rz}-e^{-rD}}{1-e^{-rD}}R(1+\frac{.05}{10}(z-NRA))$$

After time z the new objective is:

$$\max_{z,C_t} V(z) = \int_z^{z'} e^{-rt}U_W(C_t)dt + \int_{z'}^{D} e^{-rt}U_R(C_t)dt$$

s.t.

$$\int_0^z e^{-rt}C_t dt + \int_z^D C_t' dt = \int_0^{z'} e^{-rt}W_t dt + \int_{z'}^D e^{-rt}R_t dt$$

or simplifying as before, s.t.

$$C(1-e^{-rz})+C'(e^{-rz}-e^{-rD})=\left(1-e^{-rz'}\right)W+\left(e^{-rz'}-e^{-rD}\right)R(1+\frac{.05}{10}(z'-NRA'))$$

Combining the FOCs:

$$\epsilon C'=W-R(1+\frac{.05}{10}(z'-NRA'))+R\frac{.05}{10}(\frac{1}{r}-\frac{1}{r}e^{r(z'-D)})$$

$$\begin{pmatrix} 1 & \frac{-re^{-rz'}}{e^{-rz}-e^{-rD}}W+\frac{re^{-rz'}}{e^{-rz}-e^{-rD}}(1+\frac{.05}{10}(z'-NRA'))R-\frac{.05}{10}R\frac{e^{-rz'}-e^{-rD}}{e^{-rz}-e^{-rD}} \\ \epsilon & \frac{.05}{10}R(1+e^{r(z-D)}) \end{pmatrix}\begin{pmatrix} dC' \\ dz' \end{pmatrix} = \begin{pmatrix} -\frac{.05}{10}R\frac{e^{-rz'}-e^{-rD}}{e^{-rz}-e^{-rD}} \\ \frac{.05}{10}R \end{pmatrix} dNRA'$$

$$\begin{pmatrix} \frac{dC'}{dNRA'} \\ \frac{dz'}{dNRA'} \end{pmatrix}=\begin{pmatrix} 1 & \frac{-re^{-rz'}}{e^{-rz}-e^{-rD}}W+\frac{re^{-rz'}}{e^{-rz}-e^{-rD}}(1+\frac{.05}{10}(z'-NRA'))R-\frac{.05}{10}R\frac{e^{-rz'}-e^{-rD}}{e^{-rz}-e^{-rD}} \\ \epsilon & \frac{.05}{10}R(1+e^{r(z-D)}) \end{pmatrix}^{-1}\begin{pmatrix} -\frac{.05}{10}R\frac{e^{-rz}-e^{-rD}}{1-e^{-rD}} \\ \frac{.05}{10}R \end{pmatrix}$$

Solving gives that

$$\frac{dz'}{dNRA'}=\frac{\frac{.05}{10}R}{\Delta'}\left[-\epsilon\left(e^{-rz'}-e^{-rD}\right)-e^{-rz}+e^{-rD}\right]>0\ ,$$

where

$$\Delta' = .005R\left(\left(1+e^{r(z-D)}\right)\left(e^{-rD}-e^{-rz'}\right)+\epsilon e^{-rz}\left(r(z-NRA)+e^{-r(D-z)}-1\right)\right)$$
$$-\epsilon r e^{-rz}\left(W-R\right) < 0\ .$$

To show that the myopic worker has, ceteris paribus, a higher optimal age of retirement after the an increase of NRA I evaluate $\frac{dz}{dNRA}$ at $NRA' = NRA$ and $z = z'$. To show that

$$\frac{dz'}{dNRA'}(NRA' = NRA, z = z') > \frac{dz}{dNRA}.$$

after some algebra, it is sufficient to show that,

$$e^{r(z-D)}\left(1-e^{r(z-D)}\right)+r\left(\frac{R-W}{\frac{.05}{10}R}+z-NRA\right) > 0\ , \tag{10}$$

which is the same condition that determines consumption to decrease when benefits are cut.

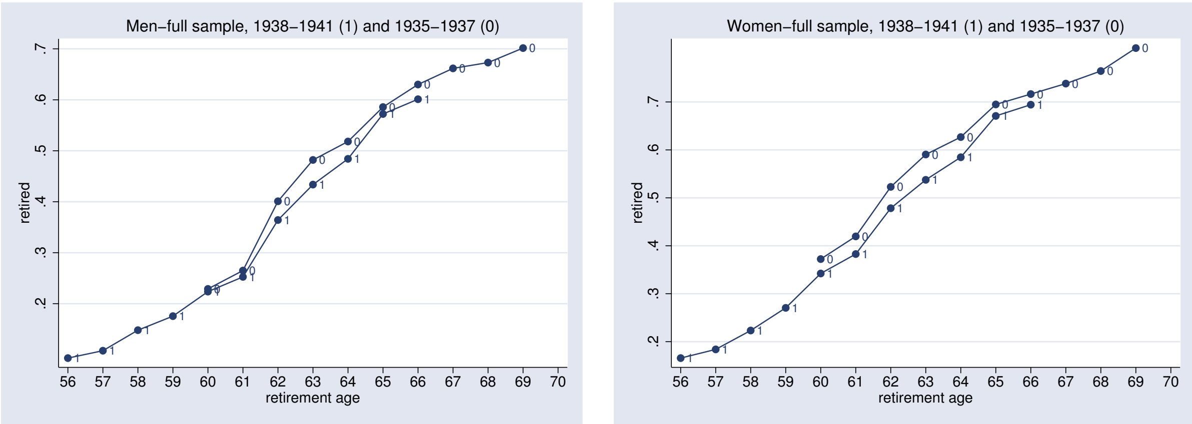

Figure 1: Cumulative distribution function of retirement age for workers affected (1) and not affected (0) by the reform.

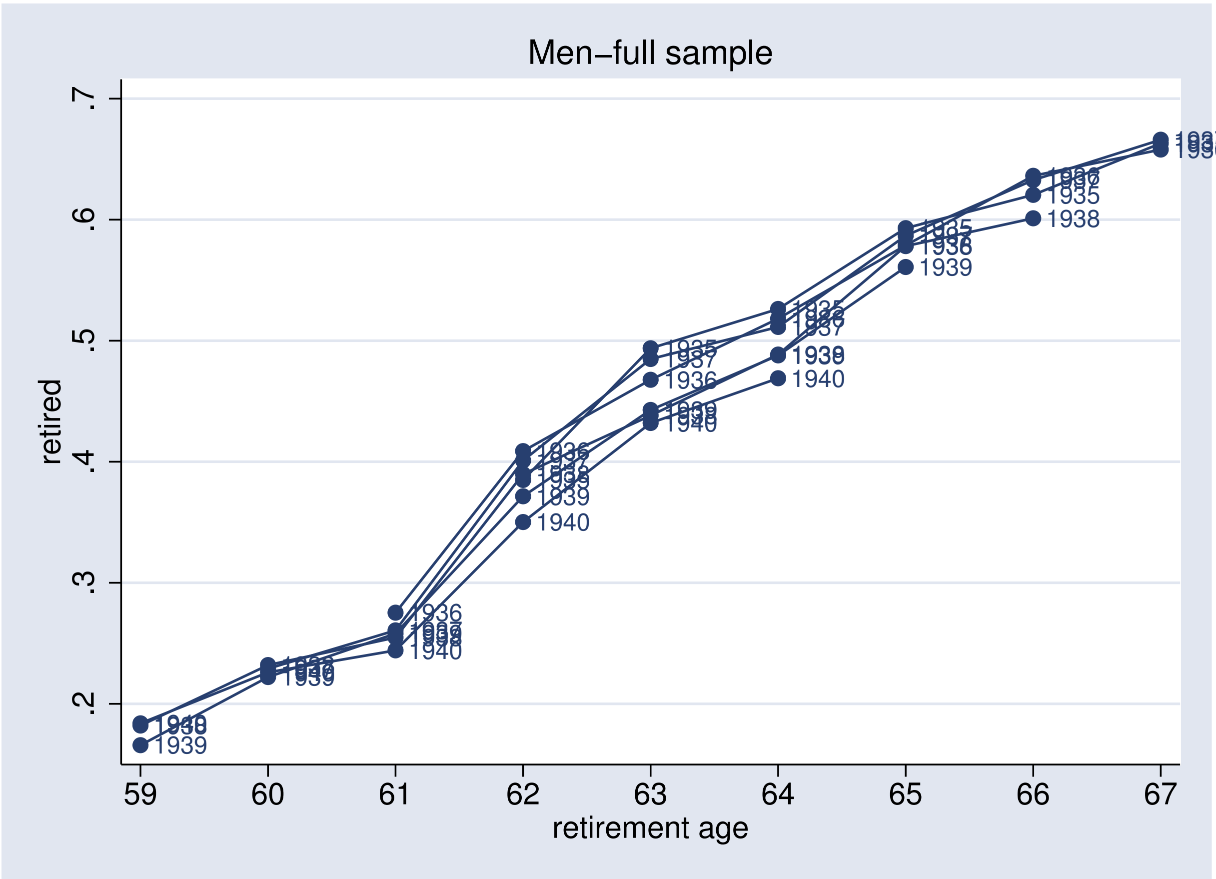

Figure 2: Cumulative distribution function of retirement age by year of birth. Full male sample.

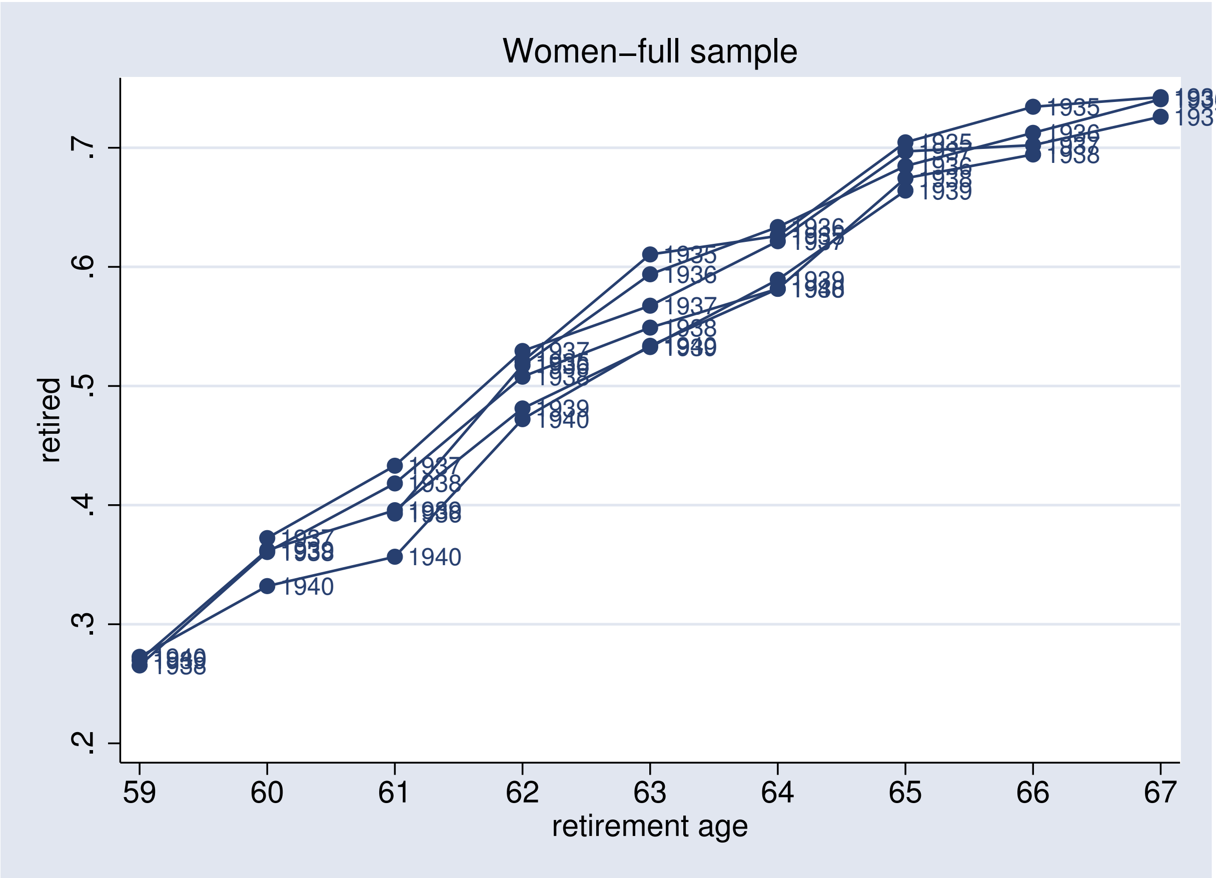

Figure 3: Cumulative distribution function of retirement age by year of birth. Full female sample.

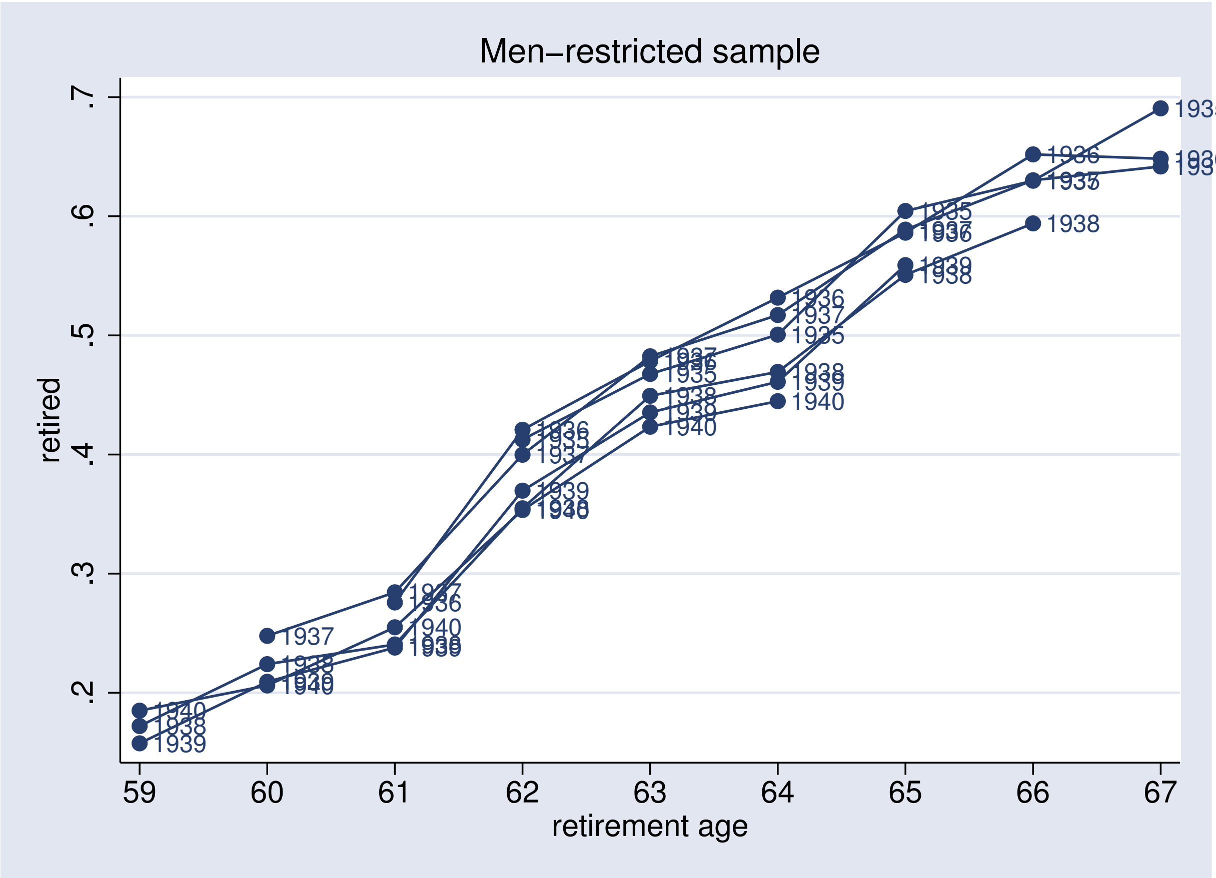

Figure 4: Cumulative distribution function of retirement age by year of birth. Restricted male sample.

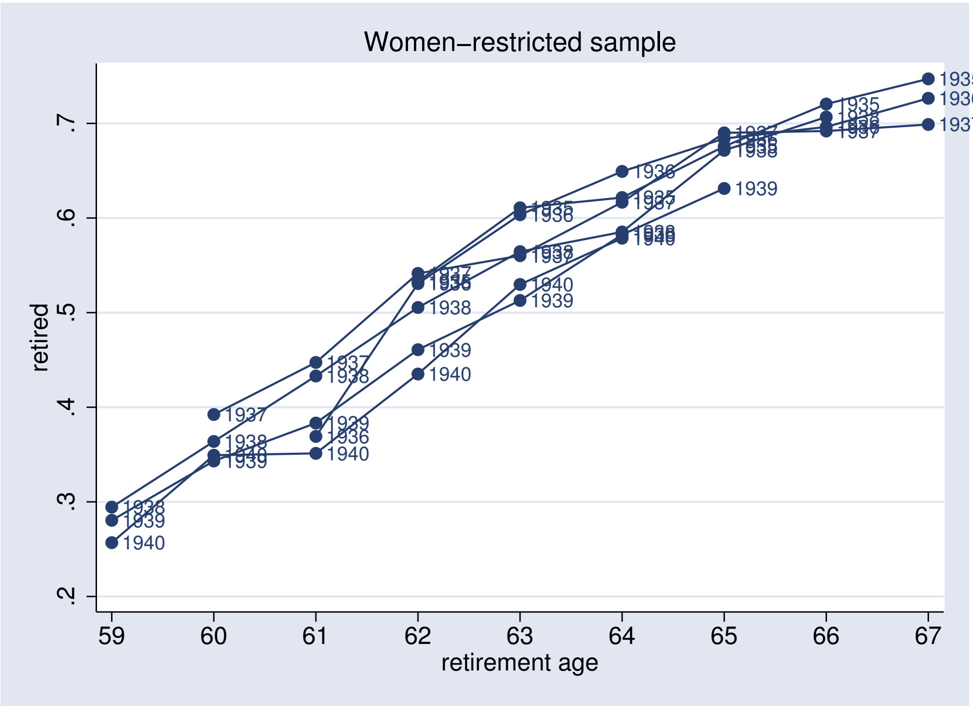

Figure 5: Cumulative distribution function of retirement age by year of birth. Restricted female sample.

Table 1: Summary statistics (mean and standard deviation) of the sample aged 61-65. Family income is reported by classes. Full sample

	1936		1937		1938		1939	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Age	63.53	1.13	63.09	1.44	63.10	1.42	63.12	1.41
Year	1999.5	1.13	2000.1	1.44	2001.1	1.43	2002.1	1.42
Male	0.49	0.50	0.49	0.50	0.49	0.50	0.49	0.50
Retired	55.58	49.69	51.67	49.97	50.15	50.00	49.47	50.00
Not Married	0.28	0.45	0.27	0.44	0.27	0.45	0.28	0.45
<High Sc.	0.19	0.39	0.18	0.38	0.17	0.37	0.16	0.36
Some college	0.15	0.36	0.16	0.37	0.16	0.37	0.17	0.37
College	0.18	0.38	0.19	0.39	0.19	0.39	0.20	0.40
Post coll.	0.09	0.29	0.09	0.29	0.10	0.30	0.10	0.31
Black	0.09	0.29	0.08	0.27	0.08	0.27	0.08	0.27
Asian	0.03	0.17	0.03	0.17	0.03	0.18	0.03	0.17
Other race	0.01	0.09	0.01	0.09	0.01	0.10	0.01	0.11
Children	0.01	0.09	0.01	0.11	0.02	0.13	0.02	0.15
#HH=1	0.17	0.38	0.16	0.37	0.17	0.37	0.16	0.37
#HH>2	0.20	0.40	0.20	0.40	0.20	0.40	0.20	0.40
Midwest	0.24	0.42	0.24	0.43	0.24	0.43	0.25	0.43
South	0.32	0.47	0.31	0.46	0.30	0.46	0.29	0.46
West	0.22	0.41	0.22	0.42	0.23	0.42	0.24	0.42
Veteran	0.25	0.43	0.24	0.42	0.23	0.42	0.22	0.41
Metrop. area	0.73	0.44	0.73	0.45	0.72	0.45	0.72	0.45
Fam.inc<20k	0.31	0.46	0.27	0.45	0.27	0.44	0.26	0.44
Fam.inc<40k	0.36	0.48	0.36	0.48	0.35	0.48	0.35	0.48
Fam.inc<60k	0.17	0.37	0.18	0.39	0.19	0.39	0.20	0.40
Fam.inc>75k	0.17	0.37	0.18	0.38	0.20	0.40	0.19	0.39

Table 2: Summary statistics (mean and standard deviation) of the sample aged 61-65. Family income is reported by classes. Restricted sample

	1936		1937		1938		1939	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Age	63.54	1.13	63.12	1.42	63.07	1.43	63.09	1.41
Year	1999.5	1.13	2000.0	1.42	2001.0	1.46	2002.0	1.44
Male	0.49	0.50	0.49	0.50	0.49	0.50	0.48	0.50
Retired	56.15	49.62	52.00	49.96	50.00	50.00	49.06	49.99
Not Married	0.28	0.45	0.28	0.45	0.28	0.45	0.29	0.45
<High Sc.	0.19	0.39	0.18	0.38	0.16	0.37	0.16	0.37
Some college	0.16	0.36	0.15	0.36	0.16	0.36	0.17	0.37
College	0.18	0.38	0.19	0.39	0.19	0.39	0.20	0.40
Post coll.	0.09	0.29	0.10	0.29	0.10	0.31	0.10	0.30
Black	0.08	0.28	0.08	0.28	0.08	0.27	0.09	0.28
Asian	0.03	0.17	0.03	0.17	0.03	0.17	0.03	0.16
Other race	0.01	0.09	0.01	0.09	0.01	0.10	0.01	0.11
Children	0.01	0.10	0.01	0.11	0.02	0.14	0.03	0.16
#HH=1	0.18	0.38	0.17	0.38	0.17	0.37	0.17	0.38
#HH>2	0.20	0.40	0.19	0.40	0.20	0.40	0.20	0.40
Midwest	0.23	0.42	0.24	0.43	0.24	0.42	0.26	0.44
South	0.32	0.47	0.31	0.46	0.31	0.46	0.29	0.45
West	0.23	0.42	0.23	0.42	0.22	0.42	0.24	0.43
Veteran	0.24	0.43	0.23	0.42	0.23	0.42	0.22	0.42
Metrop. area	0.73	0.44	0.72	0.45	0.73	0.44	0.72	0.45
Fam.inc<20k	0.31	0.46	0.29	0.45	0.27	0.45	0.27	0.44
Fam.inc<40k	0.37	0.48	0.35	0.48	0.35	0.48	0.35	0.48
Fam.inc<60k	0.16	0.36	0.18	0.38	0.19	0.39	0.20	0.40
Fam.inc>75k	0.17	0.37	0.17	0.38	0.19	0.39	0.18	0.39

Table 3: Empirical and uniform distribution of months of birth. The empirical distribution is based on 7801 certain matches born between 1937 and 1939 and aged 61 to 65.

Month	Emprical	Empirical CDF	Uniform	Uniform CDF
1	9.28	9.28	8.33	8.33
2	8.17	17.45	8.33	16.67
3	8.72	26.16	8.33	25.00
4	8.51	34.68	8.33	33.33
5	7.97	42.65	8.33	41.67
6	8.28	50.93	8.33	50.00
7	9.14	60.07	8.33	58.33
8	9.79	69.86	8.33	66.67
9	8.26	78.12	8.33	75.00
10	7.56	85.68	8.33	83.33
11	8.27	93.95	8.33	91.67
12	6.05	100	8.33	100.00

Table 4: Estimated differences in the CDFs of retirement age for the female sample.

Model	Sophisticated		Naive		Restricted	
	(1)	(2)	(3)	(4)	(5)	(6)
Age 61&Coh.36	-6.8	-7.4	-4.0	-4.6	-7.8	-8.7
	(2.8)**	(2.8)***	(1.8)**	(1.7)***	(3.0)***	(3.0)***
Age 62&Coh.36	-1.1	-1.9	-1.2	-1.9	-1.1	-1.7
	(1.9)	(1.8)	(1.5)	(1.5)	(2.6)	(2.6)
Age 63&Coh.36	4.6	3.7	2.6	2.1	4.3	3.4
	(1.8)**	(1.8)**	(1.5)*	(1.5)	(2.6)*	(2.6)
Age 64&Coh.36	1.5	1.1	1.2	0.6	3.3	2.3
	(1.7)	(1.7)	(1.4)	(1.4)	(2.5)	(2.4)
Age 65&Coh.36	-1.7	-1.7	-1.2	-1.0	-0.6	-1.1
	(1.6)	(1.5)	(1.3)	(1.2)	(2.3)	(2.2)
Age 61&Coh.38	-2.3	-1.9	-1.5	-1.1	-1.4	-0.9
	(2.2)	(2.2)	(1.4)	(1.4)	(2.5)	(2.5)
Age 62&Coh.38	-2.5	-2.1	-2.2	-1.6	-3.6	-3.7
	(2.2)	(2.2)	(1.5)	(1.4)	(2.6)	(2.5)
Age 63&Coh.38	-0.9	-0.8	-1.8	-1.3	0.4	-0.2
	(2.2)	(2.1)	(1.4)	(1.4)	(2.6)	(2.5)
Age 64&Coh.38	-4.8	-4.5	-4.0	-3.7	-3.1	-3.4
	(2.0)**	(2.0)**	(1.3)***	(1.3)***	(2.4)	(2.4)
Age 65&Coh.38	-3.0	-3.1	-2.3	-2.1	-1.8	-2.1
	(1.9)	(1.9)*	(1.2)*	(1.2)*	(2.3)	(2.2)
Age 61&Coh.39	-5.6	-5.0	-3.6	-3.3	-4.3	-3.2
	(2.3)**	(2.3)**	(1.8)**	(1.8)*	(2.9)	(2.9)
Age 62&Coh.39	-5.5	-4.3	-5.0	-3.8	-6.8	-6.4
	(2.1)***	(2.1)**	(1.7)***	(1.7)**	(2.8)**	(2.7)**
Age 63&Coh.39	-2.1	-0.8	-2.3	-0.8	-4.0	-3.6
	(2.1)	(2.1)	(1.7)	(1.7)	(2.8)	(2.7)
Age 64&Coh.39	-1.3	-0.9	-2.4	-1.8	-3.5	-3.6
	(2.0)	(2.0)	(1.6)	(1.6)	(2.6)	(2.6)
Age 65&Coh.39	-3.8	-3.1	-2.9	-2.4	-4.7	-4.2
	(1.9)**	(1.9)	(1.6)*	(1.5)	(2.5)*	(2.4)*
Other Xs	no	yes	no	yes	no	yes
Observations	103878	103878	94017	94017	18814	18814
R-squared	0.59	0.60	0.58	0.60	0.58	0.60

Notes: Robust standard errors in parentheses, * significant at 10 percent; ** significant at 5 percent, *** significant at 1 percent. Other Xs include marital status, education, race, children in the household, total members of the household, geographic region, veteran status and whether the household resides in a metropolitan area.

Table 5: Estimated differences in the CDFs of retirement age for the male sample.

Model	Sophisticated		Naive		Restricted	
	(1)	(2)	(3)	(4)	(5)	(6)
Age 61&Coh.36	-0.9	-0.8	1.5	1.5	-0.9	-0.4
	(2.5)	(2.5)	(1.6)	(1.6)	(2.8)	(2.8)
Age 62&Coh.36	-1.6	-2.1	0.8	0.3	2.1	1.4
	(1.7)	(1.7)	(1.4)	(1.4)	(2.5)	(2.5)
Age 63&Coh.36	-1.6	-2.4	-1.7	-2.2	-0.4	-0.7
	(1.9)	(1.9)	(1.5)	(1.5)	(2.7)	(2.6)
Age 64&Coh.36	0.4	-0.2	0.7	0.2	1.5	0.7
	(1.8)	(1.8)	(1.5)	(1.5)	(2.7)	(2.6)
Age 65&Coh.36	-0.2	-0.7	-0.8	-1.1	-0.3	-1.0
	(1.7)	(1.6)	(1.3)	(1.3)	(2.4)	(2.4)
Age 61&Coh.38	-1.8	-1.1	-0.6	0.0	-4.4	-4.2
	(2.0)	(1.9)	(1.2)	(1.2)	(2.2)*	(2.2)*
Age 62&Coh.38	-2.9	-2.4	-1.1	-0.7	-4.5	-4.7
	(2.2)	(2.1)	(1.5)	(1.4)	(2.5)*	(2.5)*
Age 63&Coh.38	-5.9	-5.9	-4.7	-4.5	-3.3	-3.0
	(2.2)***	(2.2)***	(1.5)***	(1.5)***	(2.6)	(2.6)
Age 64&Coh.38	-2.8	-3.0	-2.3	-2.3	-4.8	-4.8
	(2.1)	(2.1)	(1.4)	(1.4)	(2.5)*	(2.5)*
Age 65&Coh.38	-1.9	-1.8	-0.9	-0.8	-3.8	-3.7
	(2.0)	(1.9)	(1.3)	(1.2)	(2.4)	(2.3)
Age 61&Coh.39	-0.3	0.2	0.2	0.6	-3.2	-2.7
	(2.0)	(2.0)	(1.6)	(1.6)	(2.5)	(2.5)
Age 62&Coh.39	-4.0	-3.6	-2.4	-2.0	-2.8	-2.7
	(2.1)*	(2.1)*	(1.7)	(1.6)	(2.7)	(2.7)
Age 63&Coh.39	-4.9	-4.9	-4.9	-4.7	-3.1	-3.0
	(2.2)**	(2.2)**	(1.8)***	(1.8)***	(3.0)	(2.9)
Age 64&Coh.39	-1.6	-1.9	-1.7	-2.0	-5.3	-5.3
	(2.1)	(2.1)	(1.7)	(1.7)	(2.8)*	(2.8)*
Age 65&Coh.39	-1.5	-1.5	-2.4	-2.3	-2.0	-2.4
	(2.0)	(2.0)	(1.6)	(1.6)	(2.6)	(2.6)
Other Xs	no	yes	no	yes	no	yes
Observations	99688	99688	90393	90393	17823	17823
R-squared	0.48	0.49	0.48	0.49	0.47	0.49

Notes: Robust standard errors in parentheses, * significant at 10 percent; ** significant at 5 percent, *** significant at 1 percent. Other Xs include marital status, education, race, children in the household, total members of the household, geographic region, veteran status and whether the household resides in a metropolitan area.

Table 6: Estimated effect on the average retirement age (in percent of a year).

	(1)	(2)	(3)	(4)	(5)	(6)
Model	Sophisticated		Naive		Restricted	
Women						
1937-1936	3.23	1.10	1.40	-0.10	5.91	2.96
	(4.22)	(4.70)	(3.27)	(3.77)	(5.93)	(5.70)
1937-1938	-11.18	-10.51	-10.25	-8.83	-8.16	-9.43
	(4.97)**	(4.86)**	(3.86)***	(3.31)***	(5.84)	(5.82)
1937-1939	-12.71	-9.04	-12.67	-8.91	-19.06	-17.87
	(4.81)***	(4.12)**	(3.39)***	(3.19)***	(6.12)***	(5.97)***
Men						
1937- 1936	-3.08	-5.37	-1.05	-2.73	2.90	0.35
	(4.26)	(4.95)	(3.90)	(3.36)	(6.33)	(5.93)
1937- 1938	-13.55	-13.13	-8.93	-8.33	-16.38	-16.15
	(5.05)***	(4.18)***	(3.42)***	(3.85)**	(6.03)***	(5.81)***
1937- 1939	-11.91	-11.89	-11.39	-11.13	-13.16	-13.48
	(4.89)**	(4.82)**	(3.34)***	(3.28)***	(5.92)**	(6.22)**
Other Xs	no	yes	no	yes	no	yes

Notes: Sum of the coefficients of a given cohort excluding age 61. Other Xs include marital status, education, race, children in the household, total members of the household, geographic region, veteran status and whether the household resides in a metropolitan area. Robust standard errors in parentheses, * significant at 10 percent; ** significant at 5 percent, *** significant at 1 percent