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

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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 after. These cohorts began reaching retirement age in 2000. 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 about half as much as the increase in the NRA. If older workers continue to increase their labor supply, there will be 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.

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1 Introduction

In 1983, the U.S. Congress implemented an increase in the normal retirement age (NRA) of 2 months per year. Each 2-month increase in the NRA translates into a little more than a 1 percentage point reduction in Social Security benefits. This reform is likely to influence two important decisions that workers face at the end of their careers: (1) when to start collecting Social Security benefits, and (2) when to retire. Since benefits are adjusted actuarially with respect to the entitlement age, the long-term solvency of the Social Security trust fund depends more on retirement decisions than on claiming decisions. An increase in labor force participation generates more contributions, the trust fund's main source of revenue.

In this paper, I study the effects of an increase in the NRA on recent retirement behavior, providing the first ex-post evaluation of the reform. The evaluation yields both substantive evidence to guide further reforms and a guide to the calibration of structural models of retirement decisions. My results also raise serious questions about how best to improve the models on which earlier research was based. Using the change in the NRA to estimate the effect of Social Security incentives on labor supply provides additional benefits: the exact change in benefits is known, it is not prone to measurement error, and it is exogenous.

Due to the timing of the reform, I treat workers born before 1938 as the control group and workers born in or after 1938, those who experience a reduction in benefits, as the treatment group. The analysis uses monthly CPS data ranging from January 1989 to January 2006. Figure 1 shows the cumulative distribution function (CDF) of retirement age by year of birth groups. The CDF for the treated cohorts is truncated at age 67, which corresponds to year 2005 for the first treated cohort (1938). Across all birth cohorts, male workers exhibit very similar retirement patterns before age 62. For female workers, however, there is a clear trend toward later retirement at all ages.

The only age range for which the pattern of retirement of the treated cohorts differs systematically from that of the control group is between age 62 and age 65. Between these ages, treated workers (group 4), are more likely to be in the labor force than are untreated workers (groups 1, 2 and 3). This difference is even more pronounced in Figure 2, where I use a restricted sample to correct for measurement error in the year of birth variable.

Figure 3 translates the changes in the CDFs into changes in average retirement age with respect to the 1937 cohort. Because of censoring, I focus on workers younger than 66, leaving three treated cohorts: 1938, 1939, and 1940. The dotted lines show piecewise-linear fits. In all plots there is a clear break in the trend toward later retirement between the 1937 and the 1938 birth year, and the break is even more evident when I use the restricted sample.

The most obvious cause of this change in trends is the increase in the NRA. Point estimates imply an increase in the actual age of retirement of about 50 percent of the increase in the NRA for both men and women.

My estimates are at least more than three times as large as previous out-of-sample predictions. These previous predictions suggest 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 (n.d.) simulate the effect on retirement of a 1 year increase in the NRA. Depending on the specification used, the age of retirement increases between 0.5 and 2 months (using the 61-65 age range). In a similar exercise, a report for the SSA written by Panis, Hurd, Loughran, Zissimopoulos, Haider and St.Clair (2002) suggest that the average retirement age increases by about seven days. Both studies rely on estimates based on the cross-sectional variation in labor supply that is related to differences in Social Security benefits. However, Social Security benefits are endogenous because they depend on the whole history of wages. Additionally, present discounted values of future streams of benefits are likely to be measured with error, which biases the estimates downward.

Finally, the simulations only account for the financial implications of the increase in the NRA, and not for any "norms" related to the NRA (i.e., the use of the NRA as a focal point Lumsdaine, Stock and Wise 1995). Axtell and Epstein (1999), for example, suggest that the spike in the distribution of retirement age at 65 may not entirely be the product of fully rational decision-making and may instead be the outcome of herd behavior. My estimates avoid these problems. They are based on an exact and exogenous change in Social Security benefits (and their present discounted value), and incorporate changes that might be related to norms.

Despite the 1983 reform, 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, this estimate is imprecise. One of the most important sources of uncertainty is the behavior of future workers and retirees.[1] The NRA is scheduled to reach age 66, stay at that level for 10 years, and later resume the increase until it reaches age 67. Workers born in 1938. To make better predictions, it is important to keep monitoring the birth cohorts' average retirement age.

Section **??** introduces a simple intertemporal model of retirement. Its main purpose is to highlight that transitional effects arising from unexpected benefit cuts can generate large changes in the labor supply. Section 3 presents the model used for the empirical estimation. Results are shown in section 4, while section 5 concludes the paper. Appendix A describes the data.

2 A Simple Intertemporal Model of Retirement

Life-cycle theory predicts that a worker's reaction to benefit cuts-a decrease in lifetime income-will depend on when one first learns about the reform. Attentive workers may have started reacting to the reform in 1983, and after 20 years of consumption-smoothing,

[1] See Anderson, Lee and Tuljapurkar, (2003)

the change in retirement behavior is likely to be small for them. Some workers may have learned about the increase in the NRA in 1995 when the SSA began 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. Also, in 2000, the SSA added a special insert to the statement containing the changes in the NRA. Mastrobuoni (2006a) shows that the statements significantly improve workers' knowledge about their benefits. Very distracted workers may learn about the benefit at the time they claim the benefits.

The purpose of the proposed model is to show that the reaction in terms of both consumption and retirement depends on the date at which the worker learns about the benefit cut. The model is standard;it assumes that workers maximize their utility over consumption (C) and the time of retirement (z). Retirement is an 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:

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

s.t.

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

where D is the date of death. To obtain closed-form solutions, I assume that the utility function is logarithmic. 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 is the factor by which the worker's consumption must be increased to generate the same level of utility in the retiree. This 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. For simplicity, I further assume that the rate of preference equals the interest rate, $\delta = r$, and that 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. This result implicitly assumes that the Social Security rules change at time zero. Second, for reasonable parameters, if the rules change when the worker is already working, the response 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

y_i is equal to 1 when the worker is retired and zero otherwise. The following linear model (using least squares) is estimated,

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

This model measures the distance between the CDFs of retirement age of workers subject to different NRAs. Retirement is defined as not in the labor force (NILF), although results based on a more precise definition are almost identical.[2] $1(A_i = a)$ 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, $\beta_{a,c}$ measures the difference at age a between cohort c's and cohort 1937's CDF of retirement age, $\widehat{\beta}_{a,c} = E[Y|C = c, a, X = 0] - E[Y|C = 1937, a, X = 0]$. I do not include continuous Xs in the regression, hence the linear probability model is completely general.

One limitation of the data is that the year of birth variable can be misclassified. CPS data contain information about the respondent's years of age at the time of the interview.[3] Misclassification errors are not uncommon in empirical research. For example, in a paper that analyzes the impact of the earnings test on laboar supply, Gruber and Orszag (2003) take the most conservative approach of deleting observations for which ambiguity exists about the earnings test regime. Krueger and Pischke (1992) warn the reader that the probability of misclassification is approximately 20 percent when using the March CPS to establish the year of birth, but they do not explicitly correct for that.

Age at the time of the survey coupled with the information of the survey year and survey month provides, at best, an imperfect measure of the year of birth. Months of birth are almost uniformly distributed (Table 3); as a result the probability of misclassifying the year of birth based on the survey month is known. If I simply generate the birth cohort as the difference between the survey year and age, in a January survey the probability of misclassifying someone's birth year is around 11/12; someone surveyed in January is likely to have been born later in the year. The probability of misclassification is 10/12 in

[2] In 1994 the Bureau of Labor Statistics added retirement status to the labor force recode variable.

[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 few observations only. To match observations over time, I use the conservative approach of first matching by the CPS identifiers (hrhhid huhhnum hurespl), race and gender. After this first step, whenever 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).

February, and, carrying out the calculation, zero in December.[4] Using this method, the probability of misclassification would on average be 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 falls in the first half of the year reduces the average probability of misclassification to one-quarter. I call this the "naive method." When I additionally restrict the sample to the January and December surveys, the probability of misclassification is only one over twelve. I call this the "restricted method."

There is an obvious trade-off between minimizing the probability of misclassification and maximizing the statistical power. To avoid this trade-off and work with the whole sample, the restricted method makes full use of the known probabilities of 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 only depend on the survey month m, $p(m) = \Pr(C^* = c - 1|C = c, m)$. $\Pr(Y = 1|C = c, m, a, X) = E[Y|C = c, m, a, X]$ represents the conditional probability of having 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 mis-measured one is not informative, I have that

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

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

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 on a birth year (the true or the observed one), workers tend to be older later in the year. Assuming that conditional on cohort C^*, the dependence on the survey month is additively separable and does not change across cohorts, $E[Y|C^* = c, m, a] = E[Y|C^* = c, a] + g(m, a)$. 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 and 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, the true probability of cohort c being retired is a function of the observed probability, and the true probability of being retired for cohort $c - 1$,

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

As starting point for the recursion I use $E[Y|C = 1927, a] = E[Y|C^* = 1928, a]$, which implies that $E[Y|C^* = 1928, a] = E[Y|C = 1928, a]$. This allows me to analyze the differences in the CDF between several pre-reform cohorts and to properly control for preexisting trends toward earlier or later retirement.

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

$$y_i = \sum_{a=61}^{65} 1(A_i = a) \left(\sum_{c \neq 1937} \gamma_{a,c} \Pr(C_i^* = c) \right) + \gamma' X_i + \epsilon_i \; , \tag{5}$$

with the initial condition $\Pr(C_i^* = 1928) \in \{0, 1\}$.

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. The only difference is that $\Pr(C_i^* = c)$ substitutes for the previous $1(C_i^* = c)$. Like before, $\beta_{a,c}$ measures the difference between the cohorts' cumulative distribution functions.

A more easily interpretable result can be obtained from the sum of the estimated coefficients. The difference between cohort c and cohort 1937 average retirement age is

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

Tables 1 and 2 contain the summary statistics of the two samples that are used later in the analysis. The cohorts are similar in terms of racial composition and household size, though younger cohorts tend to be more educated.

Finally, to estimate the difference between the post- and the pre-1937 cohort yearly trend of the average retirement age I use the following result:

$$\Delta_{T-C} = \Delta_T + \Delta_C = \frac{1}{3}\sum_{c=38}^{40} \frac{\Delta_c}{|37-c|} + \frac{1}{9}\sum_{c=28}^{36} \frac{\Delta_c}{|37-c|} \qquad (9)$$

4 Estimation Results

I estimate Eq. (3) separately for men and women. The estimated distance between the cumulative distribution functions ($\widehat{\beta}_2$) are only shown for workers born in 1936 or later (Tables 4 and 5). Each of the three models columns (1), (3) and (5) contain only age and cohort dummies, while columns (2), (4) and (6) additionally control for marital status, education, race, total members of the household, and geographic region. Controlling for these variables reduces the estimated changes by a little. The first 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, 1939, and 1940 cohorts, and on the other side, the 1937 cohort, is usually negative. This indicates that in the 61-65 age range, the CDF of the 1937 cohort lies above the CDF of the other three cohorts.

Tables 6 and 7 show for each cohort the sum of the estimated coefficients, the sample equivalent of Eq. (8). These estimates, multiplied by 12 to obtain monthly values, represent the change with respect to the 1937 cohort of the average retirement age. Although not all post-reform $\widehat{\beta}$s are significant, most of the corresponding sums are significant at the one percent level, which suggests that the increase in the NRA generates an increase in the average retirement age. On the other hand, changes before the reform tend to be smaller and not significant.

Table 8 shows the estimates of Eq. 9 (the slopes in figure 3). The preexisting trend is larger for women than for men, but can be explained by the change in socioeconomic factors. When I control for demographic characteristics, for both men and women the preexisting trend is not significantly different from zero. The trend among the treated cohorts instead is between 1 and 1.2 months (significant at the 1 percent level). Since every year the NRA is increasing by two months, the relative change is approximately 50 percent. Controlling for other X's seems to have only a small effect. Notice also that the naive method one underestimates the effect by one-half.

4.1 Alternative Explanations

The identification is based on the assumption that the observed trend-discontinuity in the average retirement age is due to the change in the NRA. Since for the treated cohorts the estimated β_2s are negative at all ages, it is unlikely that yearly shocks are driving 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. Yet in 2000 there are already notable differences between the CDF of treated cohorts and untreated cohorts.

Also, 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 the βs 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 the 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. 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 the NRA, and at a 33 percent rate between the NRA and 69. The 33 percent rate was eliminated in 2000. The benefits that are taxed away due to the earnings test are not lost, but postponed at an actuarially fair rate. Nevertheless, evidence suggests that people perceive the earnings test as a pure tax (Gruber and Orszag 2003).

If workers decide to continue working to reach the age at which they can work without being taxed, part of the change that I attribute to the NRA reform might be due to the earnings test removal. But several factors suggest that there is no confounding. First, in 2000, the oldest treated workers are only 62 years old. A confounding effect would only be possible if spillovers reach back more than three years. Second, the earnings test removal would generate a single change, not a change in the trend.

To exclude the possibility that results are driven by labor market shocks, the same equation has been estimated using weekly hours of work as the dependent variable (excluding retirees). There are no significant differences in hours of work across these cohorts. Also, the results are not driven by differences in part-time work or disability status. Excluding disabled workers, or part-time workers (those working less than 35 hours per week) from the analysis does not alter the results.

5 Conclusions

An aging population and low labor force participation rates have worsened 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 (reducing the benefits) for workers born after 1938.

I find evidence that workers reacted strongly to this increase in the NRA. The average retirement age for cohorts that are subject to increasing NRAs is increasing by about 1 month every year, or 50 percent of the increase in the NRA. To obtain an estimated change in the average retirement trend that is based on more cohorts or on a wider age interval, the analysis presented in this paper must be repeated in a few years. But given that there is intense, ongoing work to reform Social Security, conducting early analysis with limited data is important.

Previous studies, using out-of-sample predictions, have estimated much smaller effects on labor force participation. Four major factors may have biased previous estimates toward zero. First, projections cannot capture possible changes linked to norms that are related to the NRA. Evidence suggests that some workers look at the NRA as a focal point. Mastrobuoni (2006b) shows that the distribution of the age at which treated workers claim their Social Security benefits no longer spikes at age 65, but at the NRA. Next is that estimates based on these models, since benefits are a function of past earnings, may suffer from endogeneity bias. The third source of bias 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 that this can generate larger changes in the average retirement age than would otherwise be expected. The fourth 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 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 40 years. The Social Security projections are only one of several projections made by other institutions. A common feature of all projections is that they depend heavily on the way the future behavior is modeled. My results may help evaluate the importance of an increase in 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 how the trustees forecast labor force participation. The same panel explains that the method 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, hazard rates are used instead of LFPRs. 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 if changes are smooth over time. The problem is that the increase in the NRA may have introduced a break in the trend at the end of the period used by the trustees. Therefore, it might be difficult to detect, especially if age groups are merged together. According to the 2004 Trustees report, "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 modeled directly."

The Social Security Advisory Board (*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 continues increasing the labor force participation of older workers, the trustees should follow this recommendation.

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A Data

I use the CPS monthly data from January 1989 to January 2006. The CPS data contain information about the respondent's age by the end of the survey week, usually the second week of the month.[5] I restrict the data to individuals born between 1928 and 1940, aged 61 to 65. Workers who retire early need to wait at least until age 62 before claiming their benefits. Differences in retirement rates before 62 are, therefore, unlikely to be related to the increase in the NRA. However, these restrictions represent conservative choices and may underestimate the overall effect since, as will be shown later, differences in retirement rates under age 62 and above age 65 are small, the bias is likely to be small. The CPS has a much larger sample size than the Health and Retirement Survey (HRS). For each 1928-1940 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. Another advantage of the CPS data is that the data are published soon after the interviews take place. HRS data do not contain enough treated cohorts in the age range 62 to 65.

The disadvantage of these data is that there is no information on Social Security insured status. 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). The analysis uses unweighted data. Using CPS weights results are similar, but according to the Bureau of Labor Statistics weighting revisions affected the comparability of the CPS weights over time (Bowler, Ilg, Miller, Robison and Polivka 2003).

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

B The inter-temporal 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:

$$\begin{pmatrix} \frac{dC}{dNRA} \\ \frac{dz}{dNRA} \end{pmatrix}$$

$$= \frac{1}{\Delta}\begin{pmatrix} .005R\left(-1+e^{-rD}\right)\left(1+e^{-r(-z+D)}\right) & re^{-rz}(R-W)+.005Re^{-rz}(rz-rNRA+e^{r(z-D)}-1) \\ -\epsilon\left(-1+e^{-rD}\right) & \left(-1+e^{-rD}\right) \end{pmatrix}$$

$$\begin{pmatrix} -\frac{.05}{10}R\frac{e^{-rz}-e^{-rD}}{1-e^{-rD}} \\ \frac{.05}{10}R \end{pmatrix},$$

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 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,

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

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 as 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,

$$e^{r(z-D)}\left(1-e^{r(z-D)}\right)+r\left(\frac{R-W}{\frac{.05}{10}R}+z-NRA\right)>0\ , \tag{11}$$

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

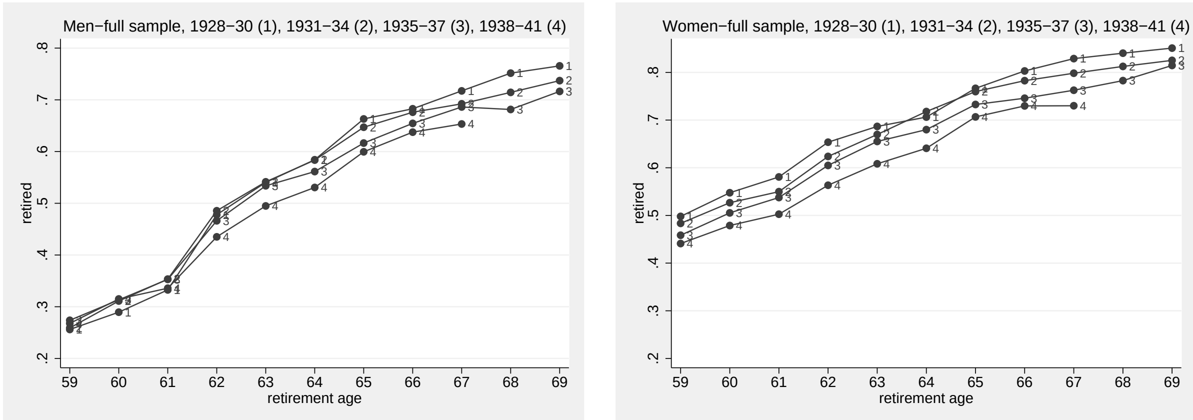

Figure 1: Cumulative distribution function of retirement age. Full sample.

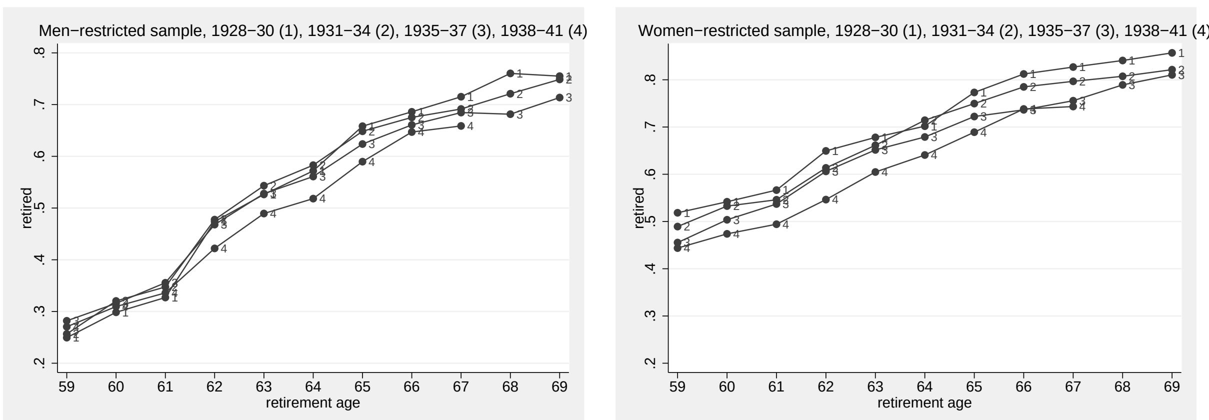

Figure 2: Cumulative distribution function of retirement age. Restricted sample.

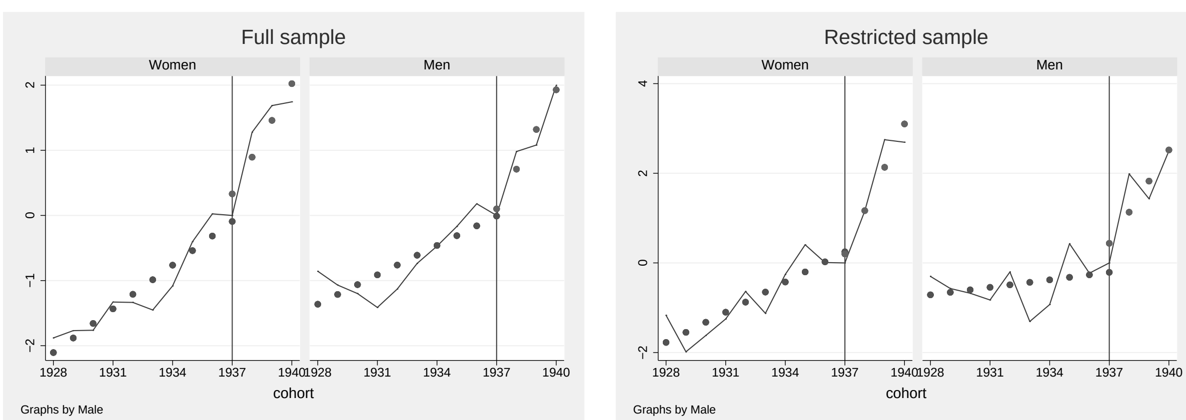

Figure 3: Change in the average retirement age (in months) with respect to the 1937 birth cohort (solid line) and its piecewise linear fit (dots).

Notes: Based on individuals between age 62 and 65.

Table 1: Summary statistics (mean and standard deviation) of the sample aged 61-65. Full sample

	1928–1930		1931–1934		1935–1937		1938–1941	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Age	62.93	1.41	62.91	1.42	63.03	1.43	63.05	1.41
Year	1992.2	1.73	1995.8	1.91	1999.5	1.74	2002.2	1.59
Male	0.46	0.50	0.47	0.50	0.48	0.50	0.47	0.50
Retired (NILF)	60.41	48.90	59.21	49.15	57.87	49.38	55.28	49.72
Employed	38.06	48.55	39.31	48.84	40.90	49.17	43.31	49.55
Not married	0.28	0.45	0.29	0.5	0.29	0.46	0.30	0.46
<High Sc.	0.27	0.44	0.25	0.43	0.21	0.41	0.18	0.39
Some college	0.14	0.35	0.14	0.35	0.15	0.36	0.16	0.36
College	0.21	0.41	0.23	0.42	0.26	0.44	0.28	0.45
Black	0.08	0.28	0.09	0.29	0.10	0.29	0.09	0.28
Asian	0.02	0.15	0.03	0.17	0.03	0.17	0.03	0.17
Other race	0.01	0.10	0.01	0.10	0.01	0.10	0.01	0.12
#HH=1	0.17	0.38	0.17	0.37	0.17	0.38	0.18	0.38
#HH>2	0.24	0.43	0.23	0.42	0.22	0.41	0.21	0.41
Midwest	0.24	0.43	0.23	0.42	0.23	0.42	0.24	0.43
South	0.31	0.46	0.33	0.47	0.33	0.47	0.31	0.46
West	0.19	0.40	0.20	0.40	0.22	0.41	0.23	0.42

Table 2: Summary statistics (mean and standard deviation) of the sample aged 61-65. Restricted sample

	1928–1930		1931–1934		1935–1937		1938–1941	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Age	62.92	1.41	62.91	1.42	63.02	1.43	63.03	1.41
Year	1992.1	1.71	1995.7	1.91	1999.4	1.74	2002.1	1.60
Male	0.46	0.50	0.47	0.50	0.47	0.50	0.47	0.50
Retired (NILF)	59.76	49.04	58.76	49.23	57.75	49.40	54.08	49.83
Employed	38.71	48.71	39.65	48.92	41.03	49.19	44.39	49.69
Not married	0.28	0.45	0.29	0.46	0.30	0.46	0.31	0.46
<High Sc.	0.27	0.44	0.25	0.43	0.21	0.41	0.19	0.39
Some college	0.14	0.35	0.14	0.35	0.15	0.36	0.16	0.37
College	0.21	0.41	0.23	0.42	0.26	0.44	0.28	0.45
Black	0.08	0.27	0.09	0.28	0.09	0.29	0.09	0.29
Asian	0.02	0.15	0.03	0.17	0.03	0.17	0.03	0.16
Other race	0.01	0.09	0.01	0.11	0.01	0.10	0.01	0.12
#HH=1	0.18	0.38	0.18	0.38	0.18	0.39	0.19	0.39
#HH>2	0.24	0.43	0.23	0.42	0.21	0.41	0.20	0.40
Midwest	0.24	0.43	0.24	0.43	0.23	0.42	0.24	0.43
South	0.31	0.46	0.32	0.47	0.33	0.47	0.31	0.46
West	0.19	0.40	0.20	0.40	0.22	0.41	0.23	0.42

Table 3: Empirical and uniform distribution of months of birth.

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

Notes: The empirical distribution is based on 7801 certain matches born between 1937 and 1939 and aged 61 to 65.

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

	(1)	(2)	(3)	(4)	(5)	(6)
	Sophisticated		Naive		Restricted	
Age 61&Coh.36	-6.5	-6.7	-3.7	-4.2	-6.9	-7.0
	(2.0)**	(1.9)**	(1.4)**	(1.3)**	(2.3)**	(2.3)**
Age 62&Coh.36	-3.7	-4.0	-1.9	-2.4	-3.4	-3.6
	(1.9)	(1.9)*	(1.3)	(1.3)	(2.3)	(2.3)
Age 63&Coh.36	1.8	1.4	2.0	1.6	2.2	1.3
	(1.9)	(1.9)	(1.3)	(1.3)	(2.3)	(2.3)
Age 64&Coh.36	1.0	0.2	0.7	0.0	1.7	0.6
	(1.8)	(1.8)	(1.3)	(1.2)	(2.2)	(2.2)
Age 65&Coh.36	-2.4	-2.4	-1.0	-1.1	-0.5	-1.1
	(1.7)	(1.6)	(1.1)	(1.1)	(2.0)	(2.0)
Age 61&Coh.38	-3.4	-3.1	-1.8	-1.6	-1.7	-1.3
	(1.9)	(1.9)	(1.3)	(1.3)	(2.2)	(2.2)
Age 62&Coh.38	-5.0	-4.7	-3.1	-2.8	-4.4	-4.4
	(1.9)**	(1.9)*	(1.3)*	(1.3)*	(2.3)	(2.2)*
Age 63&Coh.38	-2.2	-1.7	-2.0	-1.6	-0.5	-0.8
	(1.9)	(1.8)	(1.3)	(1.3)	(2.3)	(2.3)
Age 64&Coh.38	-4.2	-3.9	-3.3	-3.0	-3.0	-3.1
	(1.8)*	(1.7)*	(1.2)**	(1.2)**	(2.2)	(2.1)
Age 65&Coh.38	-3.7	-3.4	-2.2	-2.1	-1.8	-1.8
	(1.7)*	(1.6)*	(1.1)*	(1.1)	(2.0)	(2.0)
Age 61&Coh.39	-4.1	-3.6	-2.9	-2.6	-5.4	-4.7
	(1.8)*	(1.7)*	(1.4)*	(1.3)	(2.3)*	(2.3)*
Age 62&Coh.39	-6.4	-5.1	-4.9	-3.9	-9.4	-8.7
	(1.7)**	(1.6)**	(1.3)**	(1.3)**	(2.3)**	(2.2)**
Age 63&Coh.39	-3.7	-2.5	-3.4	-2.3	-4.9	-4.3
	(1.7)*	(1.6)	(1.3)*	(1.3)	(2.3)*	(2.2)
Age 64&Coh.39	-2.2	-2.0	-2.8	-2.4	-3.5	-3.2
	(1.6)	(1.5)	(1.2)*	(1.2)*	(2.2)	(2.1)
Age 65&Coh.39	-4.2	-3.5	-3.0	-2.4	-5.2	-4.7
	(1.5)**	(1.5)*	(1.2)*	(1.2)*	(2.1)*	(2.0)*
Age 61&Coh.40	-9.3	-8.3	-7.0	-6.4	-10.1	-8.7
	(2.0)**	(1.9)**	(1.5)**	(1.5)**	(2.5)**	(2.4)**
Age 62&Coh.40	-6.8	-6.0	-5.3	-4.6	-11.4	-10.6
	(2.0)**	(2.0)**	(1.6)**	(1.5)**	(2.5)**	(2.4)**
Age 63&Coh.40	-3.5	-3.1	-2.9	-2.3	-5.9	-5.4
	(2.0)	(2.0)	(1.6)	(1.5)	(2.5)*	(2.5)*
Age 64&Coh.40	-4.8	-4.0	-3.4	-2.8	-4.6	-3.7
	(1.9)*	(1.9)*	(1.5)*	(1.5)	(2.4)	(2.4)
Age 65&Coh.40	-4.4	-2.8	-3.1	-1.8	-5.4	-4.0
	(1.8)*	(1.8)	(1.4)*	(1.4)	(2.3)*	(2.3)
Other Xs	no	yes	no	yes	no	yes
Observations	440157	440157	420785	420785	84682	84682
R-squared	0.66	0.67	0.66	0.67	0.65	0.67

Notes: Standard errors clustered by individuals in parentheses, * significant at 5 percent, ** significant at 1 percent. Other Xs include marital status, education, race, total members of the household and geographic region.

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

	(1)	(2)	(3)	(4)	(5)	(6)
	Sophisticated		Naive		Restricted	
Age 61&Coh.36	0.7	0.6	2.0	2.0	0.2	0.9
	(2.0)	(1.9)	(1.3)	(1.3)	(2.4)	(2.3)
Age 62&Coh.36	-3.2	-3.4	-0.1	-0.0	0.8	1.0
	(2.0)	(1.9)	(1.3)	(1.3)	(2.4)	(2.3)
Age 63&Coh.36	-2.2	-2.2	-1.3	-1.3	-0.7	-0.4
	(2.1)	(2.1)	(1.5)	(1.4)	(2.5)	(2.5)
Age 64&Coh.36	0.6	-0.0	0.6	0.2	2.2	1.5
	(2.1)	(2.0)	(1.4)	(1.4)	(2.5)	(2.5)
Age 65&Coh.36	-1.9	-2.5	-0.8	-1.2	-0.5	-1.1
	(1.8)	(1.8)	(1.2)	(1.2)	(2.3)	(2.2)
Age 61&Coh.38	-2.9	-2.5	-1.8	-1.3	-4.4	-4.2
	(1.9)	(1.8)	(1.3)	(1.2)	(2.3)	(2.2)
Age 62&Coh.38	-4.6	-4.4	-1.5	-1.2	-6.2	-6.1
	(2.0)*	(2.0)*	(1.4)	(1.4)	(2.4)*	(2.4)*
Age 63&Coh.38	-6.4	-6.2	-4.0	-3.7	-3.4	-2.6
	(2.0)**	(2.0)**	(1.4)**	(1.4)**	(2.5)	(2.5)
Age 64&Coh.38	-2.2	-2.1	-1.8	-1.6	-3.2	-3.1
	(2.0)	(1.9)	(1.3)	(1.3)	(2.4)	(2.3)
Age 65&Coh.38	-3.0	-2.8	-0.9	-0.8	-3.7	-3.5
	(1.8)	(1.8)	(1.2)	(1.2)	(2.3)	(2.2)
Age 61&Coh.39	-0.4	-0.4	-0.0	0.1	-2.0	-2.1
	(1.7)	(1.7)	(1.3)	(1.3)	(2.3)	(2.3)
Age 62&Coh.39	-4.5	-4.2	-2.9	-2.4	-4.4	-3.9
	(1.7)**	(1.7)*	(1.4)*	(1.3)	(2.4)	(2.3)
Age 63&Coh.39	-3.6	-3.4	-3.5	-3.2	-3.5	-2.9
	(1.8)*	(1.8)	(1.4)*	(1.4)*	(2.5)	(2.5)
Age 64&Coh.39	-1.0	-1.5	-1.2	-1.5	-2.9	-3.5
	(1.7)	(1.7)	(1.4)	(1.3)	(2.4)	(2.4)
Age 65&Coh.39	-1.3	-1.6	-1.4	-1.6	-1.1	-1.7
	(1.6)	(1.6)	(1.3)	(1.3)	(2.2)	(2.2)
Age 61&Coh.40	-3.0	-2.4	-1.6	-1.0	-2.6	-2.3
	(1.9)	(1.9)	(1.5)	(1.5)	(2.5)	(2.4)
Age 62&Coh.40	-8.4	-8.0	-5.4	-4.9	-6.3	-6.0
	(2.0)**	(2.0)**	(1.5)**	(1.5)**	(2.5)*	(2.4)*
Age 63&Coh.40	-5.6	-5.2	-4.2	-3.8	-7.6	-6.8
	(2.1)**	(2.1)*	(1.6)*	(1.6)*	(2.7)**	(2.6)**
Age 64&Coh.40	-4.0	-3.9	-3.2	-3.1	-5.8	-5.4
	(2.1)	(2.1)	(1.6)*	(1.6)	(2.7)*	(2.6)*
Age 65&Coh.40	-5.7	-5.1	-3.8	-3.3	-3.8	-3.5
	(2.0)**	(2.0)**	(1.5)*	(1.5)*	(2.5)	(2.5)
Other Xs	no	yes	no	yes	no	yes
Observations	388378	388378	371779	371779	74277	74277
R-squared	0.53	0.54	0.53	0.54	0.52	0.54

Notes: Standard errors clustered by individuals in parentheses, * significant at 5 percent, ** significant at 1 percent. Other Xs include marital status, education, race, total members of the household, and geographic region.

Table 6: Estimated average retirement age (in months) minus the 1937 cohort average retirement age. Female sample.

	(1)	(2)	(3)	(4)	(5)	(6)
	Sophisticated		Naive		Restricted	
1928	2.05	1.61	1.88	1.40	1.17	0.61
	(0.39) **	(0.38) **	(0.35) **	(0.34) **	(0.59) *	(0.58)
1929	1.31	0.87	1.77	1.25	1.98	1.30
	(0.45) **	(0.44) *	(0.35) **	(0.35) **	(0.59) **	(0.58) *
1930	1.43	0.80	1.76	1.12	1.62	0.87
	(0.45) **	(0.44)	(0.35) **	(0.34) **	(0.60) **	(0.58)
1931	0.99	0.58	1.33	0.87	1.25	0.50
	(0.46) *	(0.45)	(0.36) **	(0.35) **	(0.61) *	(0.60)
1932	0.80	0.32	1.34	0.81	0.64	0.05
	(0.46)	(0.45)	(0.36) **	(0.35) *	(0.62)	(0.61)
1933	1.21	0.94	1.45	1.10	1.12	0.76
	(0.49) **	(0.48) *	(0.38) **	(0.37) **	(0.64)	(0.63)
1934	0.72	0.37	1.08	0.69	0.26	-0.10
	(0.49)	(0.48)	(0.38) **	(0.37)	(0.64)	(0.63)
1935	-0.07	-0.28	0.40	0.13	-0.41	-0.70
	(0.48)	(0.47)	(0.38)	(0.37)	(0.65)	(0.63)
1936	-0.40	-0.57	-0.02	-0.23	-0.01	-0.34
	(0.54)	(0.52)	(0.37)	(0.36)	(0.64)	(0.62)
1938	-1.82	-1.64	-1.28	-1.13	-1.16	-1.21
	(0.52) **	(0.51) **	(0.35) **	(0.34) **	(0.63)	(0.62) *
1939	-1.98	-1.58	-1.69	-1.32	-2.75	-2.51
	(0.46) **	(0.44) **	(0.36) **	(0.35) **	(0.62) **	(0.61) **
1940	-2.34	-1.91	-1.75	-1.39	-3.27	-2.85
	(0.54) **	(0.53) **	(0.42) **	(0.41) **	(0.67) **	(0.66) **
Other Xs	no	yes	no	yes	no	yes

Notes: Sum of the coefficients (times 12/100) of a given cohort excluding age 61. Other Xs include marital status, education, race, total members of the household, and geographic region. Standard errors clustered by individuals in parentheses, * significant at 5 percent, ** significant at 1 percent. The values in squared brackets represent the change in the average retirement age divided by the change in the NRA.

Table 7: Estimated average retirement age (in months) minus the 1937 cohort average retirement age. Male sample.

	(1)	(2)	(3)	(4)	(5)	(6)
	Sophisticated		Naive		Restricted	
1928	0.82	0.26	0.86	0.44	0.30	-0.07
	(0.43)	(0.42)	(0.39) *	(0.38)	(0.65)	(0.63)
1929	0.63	0.06	1.07	0.55	0.57	0.02
	(0.49)	(0.48)	(0.39) **	(0.38)	(0.66)	(0.64)
1930	0.41	-0.01	1.20	0.80	0.68	0.39
	(0.49)	(0.48)	(0.38) **	(0.37) *	(0.66)	(0.64)
1931	0.89	0.52	1.41	1.08	0.83	0.61
	(0.50)	(0.49)	(0.39) **	(0.38) **	(0.67)	(0.65)
1932	0.79	0.53	1.13	0.92	0.20	0.07
	(0.51)	(0.50)	(0.40) **	(0.39) **	(0.68)	(0.66)
1933	0.08	-0.15	0.74	0.53	1.31	1.20
	(0.53)	(0.52)	(0.41)	(0.40)	(0.71)	(0.69)
1934	0.06	-0.21	0.47	0.26	0.93	0.74
	(0.54)	(0.53)	(0.42)	(0.41)	(0.70)	(0.68)
1935	-0.37	-0.52	0.17	0.05	-0.43	-0.49
	(0.52)	(0.50)	(0.41)	(0.40)	(0.70)	(0.67)
1936	-0.79	-0.97	-0.18	-0.29	0.23	0.11
	(0.57)	(0.56)	(0.39)	(0.38)	(0.69)	(0.67)
1938	-1.95	-1.86	-0.98	-0.87	-1.99	-1.84
	(0.56) **	(0.55) **	(0.38) **	(0.37) *	(0.68) **	(0.66) **
1939	-1.25	-1.28	-1.08	-1.04	-1.43	-1.44
	(0.49) **	(0.48) **	(0.39) **	(0.38) **	(0.67) *	(0.65) *
1940	-2.85	-2.68	-1.99	-1.81	-2.83	-2.61
	(0.57) **	(0.56) **	(0.44) **	(0.43) **	(0.71) **	(0.68) **
Other Xs	no	yes	no	yes	no	yes

Notes: Sum of the coefficients (times 12/100) of a given cohort excluding age 61. Other Xs include marital status, education, race, total members of the household and geographic region. Standard errors clustered by individuals in parentheses, * significant at 5 percent, ** significant at 1 percent. The values in squared brackets represent the change in the average retirement age divided by the change in the NRA.

Table 8: Estimated trend in the average retirement age (in months).

	(1)	(2)	(3)	(4)	(5)	(6)
	Sophisticated		Naive		Restricted	
			Panel A: Female Sample			
C:1928–37	0.11	0.02	0.23	0.12	0.12	-0.01
	(0.12)	(0.12)	(0.09) **	(0.09)	(0.16)	(0.15)
T:1938–40	1.20	1.02	0.90	0.75	1.21	1.14
	(0.25) **	(0.25) **	(0.19) **	(0.18) **	(0.33) **	(0.32) **
$T - C$:	1.08	1.00	0.67	0.63	1.09	1.15
	(0.36) **	(0.35) **	(0.26) **	(0.25) **	(0.46) **	(0.45) **
			Panel B: Male Sample			
C:1928–37	-0.05	-0.12	0.12	0.06	0.11	0.06
	(0.13)	(0.12)	(0.10)	(0.09)	(0.17)	(0.16)
T:1938–40	1.17	1.13	0.73	0.66	1.21	1.14
	(0.27) **	(0.26) **	(0.20) **	(0.20) **	(0.35) **	(0.34) **
$T - C$:	1.22	1.25	0.61	0.60	1.10	1.08
	(0.38) **	(0.37) **	(0.28) *	(0.27) *	(0.49) *	(0.47) *
Other Xs	no	yes	no	yes	no	yes

Notes: Sum of the coefficients (times 12/100) of a given cohort excluding age 61. Other Xs include marital status, education, race, total members of the household, and geographic region. Standard errors clustered by individuals in parentheses, * significant at 5 percent, ** significant at 1 percent.