Consumption, Retirement and Social Security: Evaluating the Efficiency of Reform that Encourages Longer Careers*

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Abstract

This paper proposes and analyzes a Social Security reform in which individuals no longer face the OASI payroll tax after, say, age 54 or a career of 34 years, and their subsequent earnings have no bearing on their benefits. We first estimate parameters of a life–cycle model. Our specification includes non-separable preferences and possible disability. It predicts a consumption–expenditure change at retirement. We use the magnitude of the expenditure change, together with households' retirement–age decisions, to identify key structural parameters. The estimated magnitude of the change in consumption–expenditure depends importantly on the treatment of consumption by adult children of the household. Simulations indicate that the reform could increase retirement ages one year or more, equivalent variations could average more than \$4,000 per household, and income tax revenues per household could increase by more than \$14,000.

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1. Introduction

As the U.S. population ages and the moment approaches when Social Security benefit outlays will exceed payroll tax receipts, discussions of Social Security reform naturally focus on the system's solvency. Renewed concern about the Federal deficit has drawn still more attention to Social Security's assets and liabilities. This paper argues that issues of efficiency deserve greater attention as well. The current Social Security rules may generate or exacerbate labor–supply distortions; these distortions may contribute substantially to the system's social cost; and demographic trends may augment their importance. This paper proposes and evaluates a simple Social Security reform aimed at alleviating distortions to private retirement decisions that the current system may create.

The proposed reform would establish a long vesting period (say, 34-40 years of contributions). After vesting, a worker would no longer face the old–age and survivors insurance (OASI) payroll tax and his/her benefits schedule would be fixed. In fact, we would maintain the existing benefit formula, but base it only on earnings prior to the vesting age. Individuals who continue to work after vesting would thus receive a 10.6 percent payroll tax reduction. To maintain revenue neutrality within the system, there would be a small increase in the payroll tax during the vesting period.[1]

Following the tradition of Auerbach and Kotlikoff [1987] and others, we evaluate this reform in the context of a certainty equivalent life–cycle model. In contrast with that tradition, we estimate the parameters of the model using microeconomic data on earnings, consumption, and retirement. We employ what we think is a novel estimation strategy to recover key structural parameters. The strategy uses both panel data from the Health and Retirement Study (HRS) and pseudo panel consumption expenditure data from the Consumer Expenditure Survey (CEX). Simulations of the estimated model indicate that the proposed reform could raise retirement ages by more than a year, on average; equivalent variations from the reform could average $4,000 per household (2005 dollars, present value age 50) or more; and, society's additional income tax revenues could average more than $14,000 per household.

The logic of the proposed reform echoes a literature on age-dependent taxation that points to efficiency gains from using age to target lower tax rates at households with higher elasticities of labor supply.[2] Intuitively, the reform aims to eliminate the substitution effects of Social Security taxes late in life, when labor supply is especially elastic, while leaving other potential distortions of the system unchanged.

To see better how efficiency gains can arise, it helps to know that a standard assumption of our model implies that the income and substitution effects of Social Security *taxes*

[1] Similar reforms have been proposed elsewhere. See, for example, Goda *et al.* [2009] and Burtless and Quinn [2002]. This paper is, as far as we know, the first to evaluate the effects of this reform with a model with estimated parameters — see below. A related literature, including Hubbard and Judd [1987], and Hurst and Willen [2007] examines social security reforms that exempt the *young* from payroll taxes. These reforms target the inefficiencies that come from liquidity constraints that affect the young.

[2] See, for example, Kremer [2002], Erosa and Gervais [2002], Lozachmeur [2006], and Weinzierl [2010]. Banks and Diamond [2010], provide a summary.

offset one another on average. Social Security *benefits* also generate an income effect, which leads to earlier retirement, and a substitution effect, which leads to later retirement. In the case of benefits, the income effect tends to dominate; the substitution effect is slight because the present value of benefits is quite insensitive to marginal earnings for households with long work histories. On balance, therefore, the existing Social Security system tends to promote earlier retirement.

Our proposed reform eliminates the payroll tax late in careers — but before most households' optimal retirement age — canceling, for many, the tax's adverse substitution effect on work incentives. Although the positive substitution effect from the present system's benefit formula will be eliminated at the same time, its magnitude is smaller. Income effects from both taxes and benefits remain unchanged. On net, we hope to reduce work disincentives from the current Social Security system, taking advantage of the relatively high elasticity of labor supply at the age of retirement to attain significant efficiency improvements.

To quantify the effects of our reform, this paper develops a life–cycle model in which households choose their retirement age as well as their lifetime consumption/saving profile, jobs require full-time work, and retirement is permanent. The benefit to a household of later retirement is greater lifetime earnings; the cost is forgone leisure — and, more generally, lost time at home. A household derives a flow of services from its consumption expenditure and time at home. The service flow, in turn, yields utility through a conventional, concave utility function. Although our baseline model ignores health considerations, we present a second formulation with an insurable chance of disability.

The model is simple. It abstracts, among other factors, from uninsured income risk, uncertain longevity, and liquidity constraints. A benefit of simplicity is that the model offers analytic insights. One of these insights is the prediction of a discontinuous change in expenditure at a household's retirement, a change attributable to the abrupt increase in leisure and the intratemporal complementarity of expenditure and leisure.[3] A number of empirical studies have described a drop in household consumption expenditure at the time of retirement (Banks *et al.* [1998], Bernheim *et al.* [2001], Hurd and Rohwedder [2003, 2005], Haider and Stephens [2005], Aguiar and Hurst [2005], Blau [2006], and others). Our analysis shows how to use the magnitude of the drop, which this paper measures from CEX data, as well as age of retirement, measured from the HRS, to identify the model's key parameters in a simple way.[4]

We simulate our estimated model and find potentially substantial behavioral and welfare consequences from reform. We find, for example, that stopping the Social Security OASI payroll tax after a vesting period of 34 years of contributions could lead households to postpone their retirement by a year and a half or more, on average. We calculate that

[3] This non-separability may be interpreted as deriving either from tastes or from the production technology of consumption, including home production. See section 2.1.

[4] See Section 4, below, for further discussion of identification and a comparison with alternative modeling strategies. We first described our identification strategy in Laitner and Silverman [2005]. Hall [2006] develops a related method for estimating the curvature of the utility function. In a similar vein, Chetty [2006] shows how to estimate risk aversion, in part, from the change in consumption associated with a random change in labor supply.

consumers, on average, would pay as much as $4,000 (2005 dollars, in present value at age 50) to participate in the post–reform system. When we account for the social gain from income taxes on longer careers, the total social benefit could increase to more than $18-20,000 per household.

Certain assumptions of our model — such as jobs requiring full-time work, the permanence of retirement, the absence or insurability of many forms of risk, and a lack of liquidity constraints — likely amplify the behavioral consequences and efficiency gains from reform.[5] However, we believe that the estimated magnitudes of the gains in our model indicate that this paper's reform is worth further consideration.

This paper joins a large literature aimed at evaluating the effects of Social Security on labor supply. See Feldstein and Liebman [2002] for a review. By applying an explicit life–cycle model, we differ from much of this literature, which seeks reduced form estimates. Implementing a structural model allows us to evaluate the *life–cycle* effects on retirement and consumption of counterfactual reforms. By estimating the parameters of a fully-specified model, our paper also joins a smaller literature that provides structural estimates of life-cycle models of retirement (see, for example, Gustman and Steinmeier [1986], Rust and Phelan [1997], Bound *et al.* [2005], French [2005], and van der Klaauw and Wolpin [2005]). Our work is distinguished from this literature by its emphasis on a particular reform and by its use of both earnings and consumption data. Our estimation differs from many recent structural models of retirement in its certainty equivalent approach. Policy simulations, however, often employ such a framework, and we believe that it provides a rich yet tractable formulation — permitting analytic as well as numerical insights.

The organization of this paper is as follows. Section 2 describes our basic model and its formulation with stochastic disability. Section 3 discusses our pseudo–panel data on consumption expenditure, our HRS data on lifetime earnings and retirement ages, and our parameter estimates. Section 4 discusses how the model's parameters are identified, and details our estimation strategy. Section 5 qualitatively and quantitatively analyzes the Social Security reform outlined above. Section 6 concludes.

2. The Model

In this section we present the details of our basic model, and provide a sketch of how it generalizes to accommodate uncertain disability. Details of the more general model can be found in Appendix I.

2.1 Basic Model

In our model, each household maximizes utility subject to a lifetime budget constraint. We focus on married couples and assume unitary decision-making for each household and no divorce. At a household's inception, both spouses learn their earning power — that is, the lifetime profile of their wage rates. At that time, the wife sets her lifetime labor

[5] If, for example, we allowed adjustments to labor supply during the vesting period, reactions to the small increase in the payroll tax early in life would counteract efficiency gains from removing the tax late in life. Similarly, raising payroll taxes on the young would exacerbate inefficiencies that derive from liquidity constraints that we do not model.

force participation schedule, and the household chooses fertility and family structure paths. We treat these plans as predetermined when households make their remaining household consumption and labor force decisions. The real interest rate is a constant r; and life spans are certain.[6] We assume that men either work full-time in the labor market or not at all.[7] At the husband's retirement age R_i, both spouses stop working.[8] We analyze a household's choice of R_i and consumption expenditure at all ages.

We assume that the flow utility function is isoelastic. Letting x_{is} be the consumption expenditure of household i when its adult male is age s, and letting n_{is} be the household's size in "equivalent adults," the utility flow of household i at age s is

$$u(x_{it}\,,\,i\,,\,s) = \frac{1}{\gamma} \cdot n_{is} \cdot [\frac{\lambda_{is} \cdot x_{is}}{n_{is}}]^{\gamma}\,, \quad \gamma < 1\,. \tag{1}$$

The parameter λ_{is}, which captures the complementarity between time at home and consumption of goods, satisfies

$$\lambda_{is} \equiv \begin{cases} 1\,, & \text{if } s < R_i, \\ \lambda\,, & \text{if } s \geq R_i. \end{cases} \tag{2}$$

where $\lambda > 1$ is estimated below. We treat the effects of family composition as in Tobin [1967], setting

$$n_{is} = 1 + \chi^S_{is} \cdot \xi^S + \chi^K_{is} \cdot \xi^K\,. \tag{3}$$

In equation (3), χ^S_{is} equals one if household i includes a spouse at age s and equals zero otherwise, $\chi^K_{i,s}$ gives the number of "kids" that consume in the household and ξ^S and ξ^K are adult-equivalent weights (to be estimated).

[6] Some papers assume households face earnings' uncertainty (e.g., Hubbard *et al.* [1994], Gourinchas and Parker [2002], Scholz *et al.* [2006]); some assume households learn about their earning abilities (e.g., Guvenen [2007]); and some assume a nonstochastic, representative lifetime profile (e.g., Auerbach and Kotlikoff [1987]). We assume idiosyncratically different lifetime profiles, but no uncertainty about them on the part of households. We view this certainty-equivalent approach as a natural one for an initial evaluation of the proposed reform. If the reform does not generate substantial behavioral or welfare gains in the absence of uncertainty, it is unlikely that adding uncertainty will enhance the gains. In addition, allowing uncertainty would result in much less analytic tractability, a feature of the model that both clarifies identification and reveals the mechanisms behind the behavioral effects of reform. We return to the consequences of these modeling choices in Section 4, when we discuss identification.

[7] See, for example, Rust and Phelan [1997, p.786], Hurd [1996]. An indivisible workday is consistent with the fact that U.S. data show little trend in male work hours or participation rates after 1940, except for a tendency toward earlier retirement 1940-80 — e.g., Pencavel [1986], Blundell and MaCurdy [1999], and Burkhauser et al. [1999].

[8] See Gustman and Steinmeier's [2000] evidence that, in fact, couples often retire together.

To understand better the role played by the number of equivalent adults (n_{is}), temporarily set $\lambda = 1$ and $r = 0$. In that case, if we equate the marginal utility of a household's expenditure at ages s and t, we obtain

$$[n_{is}]^{1-\gamma} \cdot [c_{is}]^{\gamma-1} = [n_{it}]^{1-\gamma} \cdot [c_{it}]^{\gamma-1} \Longleftrightarrow c_{is}/c_{it} = n_{is}/n_{it} \, .$$

In this way we see that n_{is} is the scaling factor that a household uses to adjust its year–to–year expenditure in response to changes in family composition.

To understand the complementarity captured by condition (2), temporarily set $n_{is} = 1$, and think of market expenditure x_{is} as yielding a "service flow" c_{is}, upon which household utility ultimately depends. In that case our model assumes

$$c_{is} = \lambda_{is} \cdot x_{is} \, . \tag{4}$$

In other words, λ_{is} governs the household's productivity in transforming consumption expenditure into household service flows. The model allows λ_{is} to take on two different values over the life–cycle, one value while working and another (higher) value while retired.

There are several interpretations of the complementarity between consumption expenditure and leisure. (i) Upon retirement, a household no longer pays the transportation, clothing, and other costs of going to work (Cogan [1981]); it can relocate for advantages in climate, cost of living, and proximity to amenities; it gains more complete control over its schedule (e.g., Hamermesh [2005]); and, it can take maximal advantage of off–peak prices. The manifestation of these benefits in our model is $\lambda > 1$. (ii) A household has more time for leisure and/or home production after retirement (e.g., Aguiar and Hurst [2005] and House *et al.* [2008]), and non-labor-market time and x are complementary.[9]

2.1.1 The Household's Problem

We are now in a position to describe, formally, a household's problem. Starting at age S_i, household i solves:

$$\max_{R_i \, , \, x_{is}} \int_{S_i}^{R_i} e^{-\rho \cdot s} \cdot u(x_{is} \, , \, i \, , \, s) \, ds + \varphi\big(a_{i,R_i} + B_i(R_i) \cdot e^{r \cdot R_i} \, , \, i \, , \, R_i\big) \tag{5}$$

$$\text{subject to:} \quad \dot{a}_{is} = r \cdot a_{is} + y_{is} - x_{is} \, ,$$

$$a_{iS_i} = 0 \, ,$$

$$y_{is} \equiv \begin{cases} [e^M_{is} + e^F_{is}] \cdot w \cdot (1 - \tau - \tau^{ss}) \, , & \text{for } S_i \leq s < R_i, \\ 0 \, , & \text{otherwise,} \end{cases}$$

[9] In particular, specification (4) is equivalent to a model where household service flows are produced from consumption expenditure and leisure using a Cobb-Douglas technology, $c_{is} = [x_{is}]^{\alpha} \cdot [\ell_{is}]^{1-\alpha} \, , \alpha \in (0 \, , \, 1)$ where ℓ_{is} denotes leisure time.

where ρ is the subjective discount rate; the household's adult male supplies e^M_{is} "effective hours" in the labor market per hour of work; the adult female supplies e^F_{is} "effective hours;" the wage rate per effective hour is w; the income–tax rate is τ; the combined Social Security OASI and Hospital Insurance tax rate is τ^{ss}; and, household net worth is a_{is}. "Effective hours" exogenously change with age, reflecting an individual's maturity and economywide technological progress.

The function $\varphi(.)$ gives the continuation value associated with retirement at age R_i and satisfies

$$\varphi\big(A + B_i(R_i) \cdot e^{r \cdot R_i}\, , \, i \, , \, R_i\big) \equiv \max_{x_{is}} \int_{R_i}^{T} e^{-\rho \cdot s} \cdot u(x_{is}\, , \, i\, , \, s)\, ds \tag{6}$$

$$\text{subject to:} \quad \dot{a}_{is} = r \cdot a_{is} - x_{is}\, ,$$

$$a_{iR_i} = A + B_i(R_i) \cdot e^{r \cdot R_i} \quad \text{and} \quad a_{iT} \geq 0\, ,$$

where the age–0 present value of capitalized Social Security and Medicare benefits is $B_i(R_i)$. A household takes r, w, τ, τ^{ss}, e^M_{is}, e^F_{is}, and $B(.)$ as given.[10]

There may be incentives to retire at particular ages implicit in some defined benefit pension plans or employer–provided health insurance — e.g., Ippolito [1997]. Here we adopt the view that both employers and workers are heterogeneous in their preferences about retirement ages and that workers choose employers whose preferences match their own. It follows that $B(R_i)$ reflects Social Security alone — and earnings are gross of all employer benefits.

2.1.2 Optimal Consumption and Retirement

We now turn to describing some basic features of the optimal consumption path and the optimal timing of retirement. We will build on these results to derive the estimating equations below.

Proposition 1: *Let household i's problem begin at age S with optimal retirement age R. Then the solution to the household's problem satisfies*

$$x_{is}/n_{is} = (x_{iS}/n_{iS}) \cdot e^{\frac{r-\rho}{1-\gamma} \cdot (s-S)} \qquad \textit{all} \quad s < R\, , \tag{7}$$

$$x_{iR+}/n_{iR+} = [\lambda]^{\frac{\gamma}{1-\gamma}} \cdot (x_{iR-}/n_{iR-})\, , \tag{8}$$

$$x_{is}/n_{is} = (x_{iR+}/n_{iR+}) \cdot e^{\frac{r-\rho}{1-\gamma} \cdot (s-R)} \qquad \textit{all} \quad s > R\, . \tag{9}$$

[10] In our calculations, Social Security benefits begin at age R_i or age 62, whichever is larger; Medicare benefits begin at age 65; Social Security benefits depend upon retirement age and are taxed at rate $\tau/2$; and, Medicare benefits are not taxed.

Proof: See Appendix I.

The novel result is in (8), which shows how changes in expenditure at retirement are a simple function of the curvature of the utility function (γ) and the enhanced productivity of expenditure after retirement (λ).

To develop intuition for this result, note that the model assumes that the productivity of consumption expenditure rises after retirement ($\lambda > 1$). If $u(.)$ were linear ($\gamma = 1$), a household would want to increase its expenditure after retirement to take advantage of this complementarity. When $u(.)$ is concave, a second force arises: a household also wants to "smooth" its service flow over time. This creates an incentive to decrease x at retirement to offset increases in service flow c that would otherwise occur. Condition (8) reveals these competing tendencies. When utility is isoelastic, the outcome is precise: "productivity" predominates if $\gamma \in (0\,,\,1)$, and "smoothing" wins out for $\gamma < 0$.

A number of papers have documented a tendency for consumption expenditure to drop at retirement. When $\gamma < 0$, (8) predicts such a decline. Following the intuitive discussion of Proposition 1, we see that the drop is a consequence of our nonseparable utility function — and an imperfectly divisible workday.

A household's first–order condition for its optimal retirement age gives our second proposition.

Proposition 2: *Let a household's optimal retirement age be $R_i \in (S_i\,,\,T)$. Then at $R = R_i$, we have*

$$\frac{\partial u(x_{i,R-}\,,\,i\,,\,R)}{\partial x} \cdot [y_{iR} - x_{i,R-} + x_{i,R+} + B_i'(R) \cdot e^{r \cdot R}]$$
$$= u(x_{i,R+}\,,\,i\,,\,R+) - u(x_{i,R-}\,,\,i\,,\,R-)\,. \tag{10}$$

Proof: Proposition 2 is a special case of Proposition 4 below.

The first-order condition given by (10) reveals the tradeoffs that a household faces as it contemplates a slight delay in its retirement. Suppose it delayed its retirement from R to $R + dR$. The left–hand side of (10), times dR, gives the household's utility gain from delay: y_{iR} measures the gain in earnings; $x_{iR+} - x_{iR-}$ is the change in optimal expenditure corresponding to the gain in earnings; and $B_i'(R) \cdot e^{r \cdot R}$ measures incremental Social Security benefits from building a longer work history. Multiplying the sum of these dollar figures by the marginal utility of consumption converts the left–hand side of (10) to units of utility. The right–hand side, times dR, captures the household's loss in production of utility if it delays its retirement until $R + dR$, and thus forgoes leisure's beneficial effects on the productivity of expenditure.

2.2 Disability

The basic model abstracts from many forms of uninsured risk. Disability among older workers is, however, a potentially first-order impediment to reforms aimed at lengthening careers. Substantial fractions of the population may, because of poor health, find it very difficult to extend their worklives. For this reason, our quantitative investigation accommodates effects of disability on labor supply. In anticipation of that aspect of the empirical work, this section augments our basic model to include stochastic disability. For simplicity, we assume there exists actuarially fair insurance, that disability is exogenous, that one's health status is objectively verifiable, and that disability is a permanent state that prevents labor market work. We sketch the augmented model here and leave its details to Appendix I.

The complementarities between leisure and consumption expenditure remain important. In this respect, we will treat disabled households like retired households. Thus, if D_i is age of disability, we amend (2) to

$$\lambda_{is} \equiv \begin{cases} 1\,, & \text{if } s < \min\{D_i\,,\, R_i\}\,, \\ \lambda\,, & \text{if } s \geq \min\{D_i\,,\, R_i\}\,. \end{cases}$$

Disability undoubtedly lowers a household's utility. (Certainly, most people would pay to avoid it.) We could allow for a direct disutility of bad health with an additively separable term in the flow utility function. Such a term does not affect the behavior we model, however, so we omit it.

In our model, household i receives the capitalized sum $B_i(R_i) \cdot e^{r \cdot R_i}$ at its chosen retirement age whether it is disabled or not; thus, disability insurance need only tide a household over until the planned age of retirement.[11] Behavior after retirement is the same as in the basic model; hence the continuation value at the planned retirement age is unchanged.

To round out the description of a household's problem in the environment with disability, we now consider insurance. Risk aversion motivates a household to insure the earnings risk created by uncertain disability. We have assumed disability status is verifiable so, at its inception a household could choose its retirement age and sell its expected earnings stream to an insurance company for a lump sum.

However, in our model, uncertain disability generates more than just income risk; it also generates risk to household service flows since the complementarities between consumption and leisure imply that optimal expenditure at any age will depend on disability status. As a result, even complete earnings insurance will still leave the household subject to service flow risk. Thus, the household also has incentive to insure its consumption expenditure flow for ages $s < R_i$. For example, if $\gamma < 0$, a household's optimal consumption expenditure will be higher for $s < \min\{D_i\,,\, R_i\}$ than immediately thereafter.

We formalize the household's new problem, accounting for optimal disability insurance, in Appendix I. Here we describe optimal consumption in this setting with a new version of Proposition 1.

Proposition 3: *Consider the augmented model with disability. Let household i choose, at its inception, to retire at age $R = R_i$. Let the household become disabled at age $D = D_i$. Then the solution to the household's problem satisfies*

$$x_{is}/n_{is} = (x_{iS}/n_{iS}) \cdot e^{\frac{r-\rho}{1-\gamma} \cdot (s-S)} \quad all \quad s < s^* \equiv \min\{D\,,\, R\}\,, \tag{11}$$

[11] In practice, households have SSDI, worker's compensation, and possibly private disability insurance. Provided a disabled household receives SSDI, the formula for determining Social Security benefits adjusts so as not to penalize work years lost due to disability.

$$x_{is+}/n_{is+} = [\lambda]^{\frac{\gamma}{1-\gamma}} \cdot (x_{is-}/n_{is-}) \quad \mathit{for} \quad s = s^* , \tag{12}$$

$$x_{is}/n_{is} = (x_{is^*}/n_{is^*}) \cdot e^{\frac{r-\rho}{1-\gamma} \cdot (s-s^*)} \quad \mathit{all} \quad s > s^* . \tag{13}$$

Proof: See Appendix I.

The key feature of Proposition 3 is that the growth rate of consumption prior to leaving the workforce (11), the change in consumption upon leaving the workforce (12), and the growth rate of consumption after leaving the workforce (13), do not depend on whether the household stopped working due to voluntary retirement or due to disability.

The intuition for this result is as follows. Households adopt full insurance. The need to pay insurance premiums causes lifetime consumption to be lower. Given insurance, however, the onset of disability causes a household no further financial hardship. The latter fact implies that a household chooses the same proportionate expenditure change after becoming disabled as at the arrival of its planned retirement age in other circumstances.

The analog to Proposition 2 provides a first–order condition for each household's utility–maximizing retirement age:

Proposition 4: *Given an optimal retirement age $R = R_i$, if disability does not occur prior to R_i, then at $R = R_i$ we have*

$$\begin{aligned} &\frac{\partial u(x_{i,R-}\,,\, i\,,\, R-)}{\partial x} \cdot \left[B_i'(R) \cdot e^{r \cdot R} + P(R) \cdot [y_{iR} - x_{i,R-} + x_{i,R+}]\right] \\ &\quad = P(R) \cdot [u(x_{i,R+}\,,\, i\,,\, R+) - u(x_{i,R-}\,,\, i\,,\, R-)] \, . \end{aligned} \tag{14}$$

where $P(s)$ is the probability of becoming disabled after age s.

Proof: See Appendix I.

As in Proposition 2, (14) balances lost wages (adjusted for post–career expenditure changes) and sacrificed retirement benefits against an enhanced utility function after retirement. What is new is that only earnings net of the cost of disability insurance now constitute an advantage for postponing retirement and that a household's ability to generate a greater service flow from a given expenditure only rises at retirement if the household is not already disabled.

The assumption of isoelastic preferences allows us to simplify (14) for use in estimation.

Corollary to Proposition 4: *Suppose the solution to the household's problem gives optimal retirement age R_i and suppose n_{is} is continuous at age R_i. Let $\Lambda \equiv [\lambda]^{\frac{\gamma}{1-\gamma}}$. Then if disability does not occur prior to R_i, at $R = R_i$ we have*

$$\frac{B_i'(R) \cdot e^{r \cdot R} + P(R) \cdot y_{R-}}{x_{iR-}} - \frac{1}{\gamma} \cdot P(R) \cdot (\Lambda - 1) \cdot (1 - \gamma) = 0 \, . \tag{15}$$

Proof: See Appendix I.

3. Data

As previewed in the introduction, we use both CEX data on expenditures and HRS data on retirement ages, household composition, and lifetime earnings. Our estimation employs one block of moment conditions based on the optimal consumption profile described in Proposition 3 and a second based on the optimal timing of retirement described in Proposition 4. The first block relies on CEX data; the second relies on the HRS.

3.1 CEX Data

The Consumer Expenditure Survey (CEX) provides comprehensive data on U.S. household expenditures. The CEX obtains diary information on small purchases from one set of households and it conducts quarterly interviews cataloging major purchases for a second set. The survey also collects demographic data and self-reports on the value of the respondent's house. At any given time, the sample consists of approximately 5,000 (7,000 after 1999) households, which each remain in the survey for at most 5 quarters. The survey was conducted at multi–year intervals prior to 1984, and annually thereafter. This paper employs the data for 1984-2001, using survey weights to combine diary and interview data and to aggregate the quarterly data.[12]

Laitner and Silverman [2005] compares CEX annual consumption totals with the National Income and Product Accounts. Total private expenditures predicted from the CEX fall short of NIPA amounts in every year, and the discrepancy has tended to increase over time. Assuming that the NIPA numbers are accurate, and that the relative magnitude of the survey's error does not vary systematically with age, for each year we scale each CEX consumption category, uniformly across ages, to match NIPA amounts. Appendix II lists our categories and describes three additional adjustments that we make to housing services, health care, and personal business expenditures. This paper's focus does not include business cycles, liquidity constraints (see below), or consumption of the oldest elderly. And, as stated, we utilize pseudo-panel consumption-expenditure data, which averages household expenditures within cells. Beyond housing, therefore, we do not attempt to distinguish durable purchases from service flows.

Deflating with the NIPA personal consumption deflator, we derive an average consumption expenditure amount, $\tilde{x}_{st}$, for each age s and year t. We then construct a pseudo panel of consumption changes from differences between $\tilde{x}_{s+1,t+1}$ and $\tilde{x}_{st}$. We organize the CEX data so that a household's age is the age of the wife for a married couple (and the single household head in other cases). Our baseline estimation uses the sample of households ages 25-69.[13]

Changes in household composition are important inputs of our estimation strategy. The CEX provides precise information on the fraction of households that are married. However, our CEX data on children groups them into age brackets. For more accurate

[12] This paper uses raw CEX data from the ICPSR archive, and we gratefully acknowledge the assistance of the BLS in providing "stub files" of changing category definitions.

[13] We exclude older households because our model does not stress the health and mortality issues that may importantly affect consumption expenditure among the elderly.

information on children, we therefore turn to the March Current Population Survey (CPS) 1984-2001. The CPS provides data on numbers of children in each household and their ages. For purposes of estimating the weight on "kids" in the adult equivalence equation (3), we set the number of "kids" equal to the number of children age 0-17 times 0.7 plus the number age 18-25 who remain in their parents' household, up to a maximum of 2. We set the maximum because we suspect that larger households reap substantial economies of scale.[14] The weight of 0.7 on younger children follows the literature.[15]

Section 4, below, explains how identification of the model's parameters depends on the age profile of household consumption expenditure and on family composition. More specifically, the model is estimated, in part, from changes the log of average consumption and changes in family composition, by age, as measured from the pseudo-panel. Figure 1 summarizes these data presenting the average over time (from 1984-2001) of the first differences in $ln(\tilde{x})$, and these demographic variables, by age.

[**Fig 1 here**]

Our estimation also requires information on the fraction of retired households at each age and time. CEX data on retirement is unsatisfactory because the survey only asks whether the respondent is "retired" if he or she had zero weeks of work in the prior 12 months. Again, we therefore turn to the March CPS 1984-2001. We consider a CPS household retired, whether disabled or not, if the head is over 50 years old and answers that he or she is out of the labor force at the time of the March survey for reasons other than unemployment or, in the case of a male, is not "unemployed" yet reports less than 30 hours per week of work. Figure 1 also plots first differences in the fraction of retired households, by age.

3.2 HRS Data

The HRS is our second main data source. We derive earnings profiles and retirement ages from the original HRS survey cohort, consisting of households with an age-eligible respondent (age 51-61) in 1992. A majority of households signed a waiver allowing the HRS to link to their Social Security Administration (SSA) earnings history. Each history runs 1951-1991; our HRS survey data covers even years from 1992-2002. We restrict attention to once–married couples with both spouses alive in 1992, with the husband having linked SSA earnings and remaining in the labor force until at least age 52 and with the wife having linked SSA data or reporting no market work prior to 1992.[16]

[14] We also estimate the model with a maximum of 3 and obtain quantitatively similar results (see below).

[15] See, for example, Attanasio et al. [2008]

[16] In addition, we restrict attention to households with adults that have 9-24 years of education. We consider them adults at the age equaling years of education plus 6, and we drop those reaching this age before 1951. We set age of marriage at the minimum of the reported age and age at first birth. We assume that men die at the close of age 74 and women at the close of 80. We exclude couples with more than 10 years age difference or more than 10 children.

We assume that an HRS household retires when its adult male does. In each survey wave, the HRS twice asks if each adult is retired and when retirement took place. Prior to 1992, we consider a male retired if he reports that status on either question. After 1992, a male who reports being retired and works less than 1500 hours per year, or who works less than 1500 hours and never again more than 1500 hours per year, is "retired." We exclude households that pass our criterion for retirement in one survey wave but fail to do so in a subsequent wave, or that retire before (male) age 50 or remain unretired at (male) age 70. This reduces our sample to 924.[17] In sensitivity analysis, Section 4.3 shows that alternative definitions of retirement have little effect on our parameter estimates.

Appendix III provides details on our construction of lifetime earning profiles for men and women. Since HRS earnings are net of employer benefits (including health insurance, pension contributions, and employer Social Security tax), we multiply household earnings for each year by the ratio of NIPA total compensation to NIPA wages and salaries. We derive Social Security benefits after retirement from the statutory benefit formula for 2000. We also incorporate a stream of Medicare benefits after age 65, less participant SMI cost — see Appendix III.

We consider two alternative measures of disability–induced retirement. In a sequence of questions about work status, the HRS asks respondents whether they are disabled and, if so, the year of onset. According to our "stringent" definition, a male retires because of disability if he classifies himself as disabled prior to, or within one year after, retiring. In a separate sequence of questions on health status, the HRS asks respondents whether they have any health problem that "limits their ability to perform work." According to our "broad" definition of disability–induced retirement, a male retires because of disability if he classifies himself as disabled and/or as having health problems limiting his ability to work prior to, or within one year after his retirement. We exclude men who die before retiring (30 cases).

Tables 1-3 provide summary information on our HRS sample. Table 1 presents statistics on earnings and basic demographic information. The second table summarizes information on retirement ages. It divides the sample into three types of household. (i) Voluntary retirees: The male retires by choice, within sample. (ii) Involuntary retirees: The male stops working within sample but meets one or the other of our criteria for being "disabled" at the stopping time. (iii) Continuous work within sample: At the last usable age in sample, the male remains at work. Types (ii) and (iii) are treated as "censored" observations in that the desired retirement age is later than what we observe. Because our definition of retirement is fairly conservative, there are many type (ii) and (iii) households in the data. Finally, Table 3 presents cumulative fraction of men who characterize themselves as disabled and retired. Recall that the sample is limited to men who retire after age 50 and before 70. Table 3's cumulative fraction at age t corresponds to $1 - P(t)$ from

[17] This represents approximately 22% of the 4,254 married couples in the original HRS sample. Restricting attention to the once married leaves 2,750 couples. The education and within-couple age differential restrictions leave 2,280 couples. The link to SSA data leaves us with 1,474 couples. Excluding males who retire but subsequently return to work leaves a sample of 1182. A lack of early work history, combined with very late or very early retirements account for the remainder of the drop in sample.

Table 1. Statistics for HRS Couples 1992-2002

Variable	Mean	Min	Max	Coef. Var.
Age Male Last Works	60.2832	25.0000	68.0000	0.0682
Couple: Male Age - Female Age	2.7479	-8.0000	10.0000	0.9751
Male Age Marriage	24.0242	14.0000	56.0000	0.1684
Children per Couple	2.8339	0.0000	10.0000	0.4885
Male Lifetime Earnings (000s)[a]	2799.325	260.713	17177.482	0.5516
Female Lifetime Earnings (000s)[a]	584.627	0.000	4822.912	0.8296
Social Security Benefits (000s)[b]	160.082	0.000	257.924	0.1886

Source: see text. HRS household weights. Sample size=924.
(a) Present value at male (female) age 50; gross of benefits; net of OASI, HI, and federal income taxes; 2005 dollars (PCE).
(b) Present value male age 50 household net-of-income-tax lifetime OASI benefits, 2005 dollars (PCE).

Table 2. Male Retirement Status for HRS Couples 1992-2002

Category	Stringent Definition Disability	Broad Definition Disability
Voluntary retirement	444	332
Stops work due to disability	64	185
Continuous work within sample	416	407
Total	924	924

Source: see text. Definitions of “disability” — see text.

Proposition 4.

Last, our estimation requires that we specify of the rate of return on savings, the relevant tax rates, and the Social Security benefits schedule. We assume a constant gross–of–income tax real interest rate of 5%/yr. Our real interest rate comes from a ratio of factor payments to capital over the market value of private net worth. See Laitner and Silverman [2005] for details.[18] We abstract from government transfer payments other than Social Security. Our income tax rate τ comes from government spending on goods and services less indirect taxes (already removed from profits, and implicitly absent from wages and salaries below). Dividing by national income, the average over 1952–2003 is 14.28%/year. In the calculations below, the Social Security benefit formula, including the

[18] For comparison, Auerbach and Kotlikoff [1987] use 6.7%/year, and Gokhale *et al.* [2001] use 4%/yr. and 6%/yr.

Table 3. Cumulative Probability of Male Disability: HRS Couples 1992-2002					
Age	Retired and Disabled		Age	Retired and Disabled	
	Stringent Def. Male Disability	Broad Def. Male Disability		Stringent Def. Male Disability	Broad Def. Male Disability
51	.0008	.0036	61	.0748	.1714
55	.0200	.0385	65	.1541	.4153
60	.0644	.1365	70	.3608	.7794

Source: see text. See text for definitions of "disability." See Table 2 for sample sizes.

ceiling on taxable annual earnings, follows the history of the U.S. system.

4. Identification and Estimation

4.1 Intuition for Identification

In order to simulate the consequences of the counterfactual reform, we need to estimate the model's remaining utility and technology parameters

$$\beta \equiv (\rho\,,\,\gamma\,,\,\xi^S\,,\,\xi^K\,,\,\lambda)\,. \tag{16}$$

We adopt a method of moments approach with two blocks of moment conditions, one based on Proposition 3 and CEX data, and the other on Proposition 4 and the HRS. Before presenting the details of the estimation strategy, here we provide intuition about what features of the data identify the model's parameters.

4.1.1 Intuition for the CEX Moments

Parameter estimation from CEX moments rests on equations (11)-(13). Those results show that, in the absence of family composition changes or retirement/disability, the log of a household's consumption is predicted to rise or fall at a constant rate $\frac{r-\rho}{1-\gamma}$. This is a familiar formula. It shows that, if household composition and labor supply remain unchanged, the age-profile of log consumption is dictated by a combination of time discounting (ρ, given r) and the taste for smoothing $(1-\gamma)$.

Of course, household composition and labor supply change over the life–cycle. Families form, children are born and eventually leave home, households retire or become disabled, and, finally, spouses die. Equations (11)-(13) show how these changes affect the change in log consumption.

When a child is added (or leaves the house), the number of adult equivalents rises (falls) discretely, and so does the log of consumption. The magnitude of this change in consumption is dictated, in part, by the size of the adult equivalency weight on kids ξ^K. Analogous changes in consumption occur upon the addition or loss of a spouse. Similarly, equation (12) shows that when a household retires or becomes disabled, the log of consumption may also change discretely. In this case, as discussed above, the size and direction of the change depends on the degree of complementarity between consumption and leisure (λ) and the taste for smoothing $(1-\gamma)$.

Taken together, the effects of all of these parameters lead to the following prediction: the log of consumption rises (or falls) linearly with age, except when it jumps discretely with changes in household composition or labor supply.

The preceding discussion applies to a single household. In fact the parameters in (16) are estimated from pseudo-panel data representing age-year averages taken over many households. The same intuitions apply but, now, the log of consumption expenditure will change more smoothly as *average* household size and labor supply change in smaller increments. These are precisely the changes that are displayed in Figure 1.

A brief examination of Figure 1 offers a preview of the estimation results. Figure 1 shows that the growth rate of log consumption expenditure is positive and relatively large until households reach their mid to late 40s. Before these ages, households are also adding members, both spouses and children. Therefore the model must attribute the rising average consumption to some combination of increasing household size and relative patience (ρ must be small relative to r).

As households reach their mid to late 40s, their children start leaving at a higher rate and a few households start to retire or become disabled. The model predicts that this combination will slow average consumption growth, as is consistent with the CEX data. In the next stage of life, as households reach their late 50s and early to mid 60s, they start retiring or becoming disabled in large numbers. Previous studies tend to find a decline in consumption expenditure at retirement or disability, thus we would expect these events to slow average consumption growth. The magnitude of this slowdown will depend on all of the model's parameters, as they dictate the rate of growth absent retirement and the size of the change in consumption expenditure upon retirement.

4.1.2 Intuition for the HRS Moments

As the previous discussion suggests, the CEX consumption profiles permit identification only of *composites* of the parameters in (16). In particular, the average rate of growth in consumption identifies $\frac{r-\rho}{1-\gamma}$, and the change in consumption expenditure at retirement identifies $\frac{\gamma}{1-\gamma} \cdot \ln(\lambda)$. Even with information about r, we cannot separately identify ρ, γ, and λ. Unpacking the structural parameters requires more information. For this we turn to the optimal retirement equation (15).

Intuitively, the HRS data on earnings and household composition dictate, for any choice of the parameters in (16), an optimal retirement age for each household. By pinning down $\frac{r-\rho}{1-\gamma}$, and $\frac{\gamma}{1-\gamma} \cdot \ln(\lambda)$, the CEX moments restrict the set of composite parameters that are feasible. From that set, our estimation procedure selects the parameters that best fit the HRS data on retirement ages given net-of-tax income paths and household composition. In this respect, a key intuition involves the effect of λ on the age of retirement. If λ is too low, households will never retire, if λ is too high, households will retire too young. Details of the estimation procedure follow.

4.2 Estimation

In this subsection we describe our estimation equations in detail.

4.2.1 Estimation from the CEX

Approximate the adult equivalent equation (3) with

$$\ln(n_{is}) \approx \chi_{is}^S \cdot \xi^S + \chi_{is}^K \cdot \xi^K ,$$

and let χ_{is}^{RD} denote an indicator equal one if household i is retired or disabled at age s, and zero otherwise; let χ_{is}^S equal 1 if the household has a spouse, and zero otherwise; and, let χ_{is}^K equal the number of kids (as described above). Let β^{CEX} denote the vector of composite parameters estimable from the CEX alone,

$$\beta^{CEX} \equiv \big(\frac{r-\rho}{1-\gamma} , \xi^S , \xi^K , \frac{\gamma}{1-\gamma} \cdot \ln(\lambda)\big) . \tag{17}$$

Then Proposition 3 shows

$$\Delta \ln(x_{is}) \approx \beta_1^{CEX} + \beta_2^{CEX} \cdot \Delta\chi_{is}^S + \beta_3^{CEX} \cdot \Delta\chi_{is}^K + \beta_4^{CEX} \cdot \Delta\chi_{is}^{RD} \tag{18}$$

where Δy_{is} denotes the change in household i's y between ages s and $s+1$. In short, the change in the log of a household's consumption is a linear function of the composite parameters β^{CEX}.

The CEX provides average expenditure amounts for each age s and time t ($\tilde{x}_{st}$). It also provides analogous information about average rates of marriage ($\tilde{\chi}_{st}^S$). As noted above, we calculate the relevant averages for number of children ($\tilde{\chi}_{st}^K$) and fraction retired or disabled ($\tilde{\chi}_{st}^{RD}$) from the CPS. We can thus replace the household-level equation (18) with its age-time average analog (applying sample weights) and define the CEX moments:

$$q_{st}^1(\beta^{CEX}) \equiv \Delta \ln(\tilde{x}_{st}) - \beta_1^{CEX} - \beta_2^{CEX} \cdot \Delta\tilde{\chi}_{st}^S - \beta_3^{CEX} \cdot \Delta\tilde{\chi}_{st}^K - \beta_4^{CEX} \cdot \Delta\tilde{\chi}_{st}^{RD}$$

where $\Delta\tilde{y}_{st}$ denotes the change in the household average of y between age-year (s,t) and age-year $(s+1,t+1)$.

Our estimation then uses

$$q_{st}^1(\beta^{CEX}) = \upsilon_{s+1,t+1} - \upsilon_{st} , \tag{19}$$

where υ_{st} is an iid random variable with mean 0 and variance σ_υ^2.

We think of υ_{st} as reflecting measurement error in $\ln(\tilde{x}_{st})$. A household survey of consumption expenditure places particularly heavy burdens on respondent memory and understanding of classification systems; systematic errors and omissions can affect even average values such as $\ln(\tilde{x}_{st})$.

Although we estimate (19) as a GLS regression, for symmetry with our second block we interpret the GLS first–order conditions as method of moments equations.

4.2.2 Estimation from the HRS

Our second block of moments is given by the first-order condition in Proposition 4 and uses lifetime earnings, demographic information, and retirement ages from the HRS. Given candidate parameter values for β^{CEX} and γ, HRS data for household i, and an age of last earnings $R = R_i^0$, Proposition 3 characterizes the household's optimal expenditure time path x_{is}. Using that expenditure time path and the Corollary to Proposition 4, define

$$g_i^2(\beta^{CEX}, \gamma, R) \equiv \frac{B_i'(R) \cdot e^{r \cdot R} + P(R) \cdot y_{i,R-}}{x_{i,R-}} - \frac{1}{\gamma} \cdot P(R) \cdot (e^{\beta_4^{CEX}} - 1) \cdot (1 - \gamma)\,.$$

If R_i is the optimal retirement age for the household and $R_i^0 = R_i$, then $g_i^2 = 0$.

In practice, if $R_i^0 = R_i$, we set $g_i^2 = \epsilon_i$ where ϵ_i is a random variable with mean 0. We interpret this residual as reflecting measurement error — specifically for $y_{i,R-}$. Although $B_i'(R)$ and $x_{i,R-}$ reflect average lifetime earnings, $y_{i,R-}$ only registers last earnings, which happen to be particularly difficult to assess accurately — see Appendix III.[19]

If disability (or death) causes the household to stop working prior to the optimal retirement age (i.e., if $R = R_i^0 \leq R_i$), we assume $g_i^2 \geq \epsilon_i$.[20] Similarly, if at the last wave of the survey household i has not yet retired, again $R_i^0 \leq R_i$ — so we assume $g_i^2 \geq \epsilon_i$.

Assuming that ϵ_i is normally distributed, let $\phi(., \sigma_\epsilon^2)$ be the normal density and define

$$q_i^2(\beta^{CEX}, \gamma, R) \equiv \begin{cases} g_i^2(\beta^{CEX}, \gamma, R) & \text{if voluntarily retires in sample,} \\ \dfrac{\int_{-\infty}^{g_i^2(\beta^{CEX}, \gamma, R)} e \cdot \phi(e, \sigma_\epsilon^2)\, de}{\int_{-\infty}^{g_i^2(\beta^{CEX}, \gamma, R)} \phi(e, \sigma_\epsilon^2)\, de} & \text{otherwise,} \end{cases}$$

$$q_i^3(\beta^{CEX}, \gamma, R) \equiv \begin{cases} [g_i^2(\beta^{CEX}, \gamma, R)]^2 & \text{if voluntarily retires in sample,} \\ \dfrac{\int_{-\infty}^{g_i^2(\beta^{CEX}, \gamma, R)} e^2 \cdot \phi(e, \sigma_\epsilon^2)\, de}{\int_{-\infty}^{g_i^2(\beta^{CEX}, \gamma, R)} \phi(e, \sigma_\epsilon^2)\, de} & \text{otherwise.} \end{cases}$$

Then our HRS moment conditions are

$$\sum_i q_i^2(\beta^{CEX}, \gamma, R_i^0) \cdot 1 = 0 \quad \text{and} \quad \sum_i q_i^3(\beta^{CEX}, \gamma, R_i^0) \cdot 1 = \sigma_\epsilon^2\,. \tag{20}$$

Given the composite parameters estimated from the CEX alone, $\widehat{\beta^{CEX}}$, (20) determines $(\gamma, \sigma_\epsilon^2)$. The estimate of γ, denoted $\widehat{\gamma}$, and the first and fourth elements of the composite parameter vector $\widehat{\beta}^{CEX}$ together give our estimates of ρ and λ. Thus, we have our estimate of all of the structural parameters, $\widehat{\beta}$. We estimate the covariance matrix of $\widehat{\beta}$, as in Gallant [1987, ch.6].[21]

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[19] Note that if, for example, $y_{i,R-}^*$ is the actual value but we observe $y_{i,R-} = y_{i,R-}^* \cdot (1 + \eta_i)$, $\eta_i \sim N(0, \sigma_\eta^2)$, then $\epsilon_i = (y_{i,R-}^*/x_{i,R-}) \cdot \eta_i$ will be roughly homoscedastic because age-earning profiles for different individuals tend to be roughly parallel and because in our formulation, proportionate changes in earnings leave $y_{r,R-}^*/x_{i,R-}$ and R constant.

[20] The second–order necessary condition for R_i implies g_i^2 is decreasing in R at $R = R_i$ (recall the discussion following Proposition 2). Our inequality takes this monotonicity to be global.

[21] Some analyses (e.g., Gustman and Steinmeier [2000]) argue that the random variable

4.3 Baseline Results

Table 4 presents our baseline parameter estimates. The first column in the top panel displays baseline estimates of β^{CEX} from equation (19). The bottom panel of this column presents estimates of the structural parameters β, given β^{CEX}, for the two different definitions of disability.

Note that a sensible interpretation of our model requires $\gamma < 1$, $\lambda > 1$, $\xi^S > 0$, and $\xi^K > 0$, and that these conditions hold for the baseline estimates, regardless of the definition of disability. Estimates of the preference parameters indicate curvature very near, but statistically different from, the log case ($\gamma = -0.0797$, and significantly different from zero), indicating an elasticity of intertemporal of substitution (EIS) of about 0.93. Estimated time discounting is qualitatively modest, but significantly different from zero ($\rho = 0.0143$). Finally, the estimates point to a qualitatively large degree of complementarity between consumption and leisure ($\lambda = 2.77$ to 3.08). The table's bottom panel shows that the estimates are not very sensitive to the definition of disability.

The curvature parameter can be easily compared with previous estimates and calibrations. Using aggregate consumption data Hall [1988], Campbell and Mankiw [1989], and Patterson and Pesaran [1992], for example, estimate an EIS for consumption to be very nearly zero. Micro studies tend to estimate larger intertemporal elasticities. Banks *et al.* [1998], for instance, estimate the average EIS for consumption to be approximately 0.5. In another example, Attanasio and Weber [1993] estimate an EIS for consumption of approximately 0.75 from micro data. Using data similar to ours, Gourinchas and Parker (2002) estimate an EIS of approximately 2.0. Although our calculations rely on a very different source of variation, our curvature estimates are thus similar to those obtained in the micro studies, if a bit closer to the log case.

The time discount rate is not separately identified, without an assumption on the interest rate. A simple comparision of discount rate estimates in the literature is, therefore, not particularly informative. It is, however, important to note that the time discount rate that we estimate (approximately 1.5%) is lower than the (net-of-tax) interest rate that we have calibrated (4.3%). This relative patience is different from what emerges from studies that adopt an uncertainty approach, such Gourinchas and Parker [2002] or Scholz *et al.* [2006]. Gourinchas and Parker [2002] estimate a discount rate that is about 0.005 higher than the real interest rate. Scholz *et al.* [2006] calibrate a discount rate equal to the real interest rate. As noted in Section 4.1, because we have abstracted from income uncertainty and liquidity constraints, relative patience is necessary to explain the rising consumption profile early in the life–cycle. With earnings uncertainty and liquidty constraints, the rate of consumption growth declines late in life because, then, the constraints no longer bind, buffer stock saving is no longer optimal, and consumers are relatively impatient.

Our parameter λ is less typical and thus harder to compare with previous estimates. However, a number of papers use an equivalent Cobb-Douglas formulation — see our

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ϵ_i might reflect a household's taste for time at home as well as measurement error. Our orthogonality condition only requires that ϵ_i have mean zero. Even if heterogeneity of taste were important, our estimates could remain consistent. This paper's focus is policy reform, but the nature of the block-2 error term is an interesting topic for future research.

footnote (9). Suppose that a week has $7 \times 12 = 84$ available hours and that a full–time workweek is 40 hours. Then the Cobb-Douglas formulation might imply $\lambda = [\bar{\ell}]^{(1-\alpha)/\alpha}$ where $\bar{\ell} = 84/44 = 1.91$. Auerbach and Kotlikoff's [1987] favorite calibration has $\bar{\gamma} = -3$ and α (roughly) $= 0.4$; hence, in our terminology they assume $\gamma = \alpha \cdot \bar{\gamma} = -1.2$ and $\lambda = [1.91]^{.6/.4} = 2.64$. Altig et al. [2001] use $\bar{\gamma} = -3$ and α (roughly) $= 0.5$; hence, $\gamma = -1.5$ and $\lambda = [1.91]^{.5/.5} = 1.91$. Cooley and Prescott [1995] set $\bar{\gamma} = 0$ and $\alpha = .36$; consequently, in our terminology, they have $\gamma = 0$ and $\lambda = [1.91]^{.64/.36} = 3.16$. Our estimates of γ and λ in Table 4 thus fall in the general range of these prominent calibrations.

As noted in our intuitive discussion of identification (Section 4.1), our structural estimates depend critically on the composite parameters estimated just from the CEX moments, $\widehat{\beta}^{CEX}$. Our baseline estimate of $\widehat{\beta}_1^{CEX}$ indicates an average lifetime growth rate for households' per capita consumption (see Proposition 3) of 2.6%/yr. The estimates of the adult equivalency weights, ξ^S and ξ^K, are consistent with substantial returns to scale for larger households. The estimate of ξ^S suggests that the addition of a spouse raises household consumption by 34 percent. We estimate a nearly identical increase associated with each child age 18-25 who remains in school or in his/her parents' household (times 0.7 for younger children). For comparison, Mariger [1986] estimates that children consume 30 percent as much as adults; Attanasio and Browning [1995, p. 1122] suggest 58 percent; Gokhale *et al.* [2001] assume 40 percent.

The baseline estimate of $\widehat{\beta}_4^{CEX}$ in Table 4 indicates an 8.3 percent drop in consumption at retirement. This falls roughly in the middle of the range of estimates in what is now a very large literature on this subject. See Hurst [2008] for a review. In a previous version of this paper we estimated a considerably larger decline in expenditure at retirement. As we show in subsection 4.4.2, that result was driven by our prior treatment of the consumption expenditure of adult children who are living with their parents or in school.

4.4 Sensitivity and Goodness of Fit

This subsection evaluates the consequences for the parameter estimates of changing various modeling assumptions, and it assesses the model's in-sample fit.

4.4.1 Sensitivity

We begin by investigating the effects of relaxing some of our assumptions regarding uncertainty. The nature of our data precludes analysis of idiosyncratic shocks to individual households. The coefficients on annual time dummy variables can, however, capture aggregative shocks to interest and/or wage rates. The second column of Table 4 evaluates the effects on parameter estimates of controlling for such shocks. Comparing the results across columns 1 and 2 of Table 4, we see that adding annual time dummy variables changes our parameter estimates only modestly. Similarly, adding controls for annual birth cohort dummies in the HRS sample (dropping the very small 1926 and 1942 cohorts), has only a minor effect on parameter estimates. See the first column of Appendix Table 4B.

The Euler equation literature interpretes the error of equation (19) in terms of rational expectations. Simple time dummies will capture this only imperfectly as earnings and other shocks may have age-dependent implications. The second column of Appendix Table 4B provides estimates after averaging over each age for the entire CEX sample, and omitting time dummies. There is a drastic decline in degrees of freedom, of course, and standard errors rise accordingly. Qualitatively, however, parameter estimates are similar to those from our baseline specification. Most notable, in this specification the point estimate of the equivalency weight for a spouse ξ^S increases to 0.526, and we detect virtually no decline in consumption expenditure at retirement.

As discussed in Section 4.1, household consumption tends to rise in early adulthood and our modeling strategy attributes this rise to patience (ρ small relative to r) along with the consolidation of singles into couples and the birth of children. Other models explain the rise with relative *impatience* while adding liquidity constraints or learning about earnings potential.[22] Restricting the CEX sample to later and later ages, we look for evidence that our model is mis-specificied in this regard. The results are presented in the first columns of Appendix Tables 4D and E. We find little evidence that, by abstracting from liquidity constraints or learning about life-time income, we are improperly attributing the rise in consumption with age to relative patience. Indeed, when we raise the youngest age to 40 when liquidity constraints are presumably relaxed and more about life-time income is known, the estimated growth rates of expenditure with age ($\widehat{\beta_1^{CEX}}$), *increases* and our estimate of ρ declines.[23] We interpret these results as indicating that the model does not require liquidity constraints or learning about life-time income in order to adequately explain these data.

Another natural concern is that, since we impose a relatively conservative definition of retirement, many of our households have censored retirement choices. Our estimation strategy deals with this censoring, in the spirit of a Tobit model, by assuming a parametric form for the residuals that will determine unexplained heterogeneity in censored retirement dates. (See Section 4.2). In the second column of Appendix Table 4C we estimate our baseline model with a sample that drops all censored observations. The parameter estimates are very little changed. In this way, we find no evidence that our method of dealing with the censoring is importantly driving the results.

4.4.2 The Importance of Older Children's Consumption

In an earlier version of this paper, we included in our measure of the number of children in a household only those of age 0-17. As noted above, we have since adjusted our measure to include children age 18-25 who are still living at home or attending school full-time. Here we show the consequences for parameter estimates of these two treatments of older children.

The top panel of the second column of Appendix Table 4F presents results where we adopt the treatment of adult children from the previous version of this paper — and include CEX respondents up to age 79, as in the previous work. Relative to our preferred specification, the estimated decline in consumption expenditure at retirement/disability is much larger (36%), and the estimates of the child weight and time discount rate are

[22] Early examples include Mariger [1986], Zeldes [1989], and Hubbard and Judd [1986]. More recent models also incorporate stochastic earnings, – e.g., Deaton [1991], Hubbard et al. [1995], Gourinchas and Parker [2002], and Guvenen [2007].

[23] In a similar spirit, we can also raise or lower the oldest age in the CEX sample with little effect on the parameter estimates. See Appendix Tables 4E and F.

smaller (20% and 0.5-0.6%, respectively). The first column of Table 4G shows that the differences in these estimates are not due to the differences in the ages included in the CEX sample. Instead, the significant changes stem from the child regressor.[24,25]

Why does the treatment of older children importantly affect estimates of the change in expenditure upon retirement or disability? In the previous version of the paper, household expenditure on children age 18 and older was interpreted as their parents' consumption. The "departure" of these children at age 18 thus appeared to have a relatively modest effect on total household expenditure. The estimated child weight (ξ^K) was, correspondingly, smaller. Now recall the intuition developed in section 4.1. Other things equal, a smaller child weight implies a decrease in the estimated rate of time discount, because the arrival of children must now have a smaller effect on household expenditure, yet the rise in average expenditure during early adult years must still be explained. At the same time, the model needs to explain the slow-down in average consumption growth for household in their late 40s to early 60s. Children depart the household in these times of life but, as just noted, the smaller child weight implies that their departure must cause a relatively modest decline in household expenditure. The model must therefore attribute the slowdown in average household expenditure growth to changes household labor supply. The result is a relatively large estimate of the decline in expenditure upon retirement or disability.

We believe that the new formulation better reflects household composition, including the costs of school and other support for young adult children. Interestingly, despite important differences in the parameter estimates that emerge in this new formulation, the simulated consequences of the reform are qualitiatively similar.

4.4.3 Goodness of Fit

Turning to goodness of fit of the consumption data, the model appears to capture the life-cycle pattern of expenditure fairly well. For the specification with time dummies, Figure 2 plots residuals from the left–hand side of (19), averaged over $t = 1984, ..., 2000$, for each age s. As one would expect from such a simple model, the residuals are sometimes large in magnitude. However, we see little evidence that the model systematically mispredicts changes in log consumption at particular points in the life-cycle.

[Fig 2 here]

Since we do not base our HRS estimates on a conventional regression equation, we have no exact analog to Figure 2 for evaluating goodness of fit; nevertheless, Figure 3 may serve as a substitute. Figure 3 presents the cumulative distribution of retirement ages for couples that voluntarily retire in the sample of Table 4, column 1. We present one panel for each quartile of the distribution of the (at age 50) net present value of household

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[24] The assumption that households have at most two children does not seem to affect the estimates importantly. We see little change when we increase the maximum number to 3. See Appendix Table 4G, column 2.

[25] Even Table 4F, column 2, does not exactly duplicate the earlier version of the paper. For example, Table 4F weights children 0-17 with 0.7 (rather than 1.0) and measures children from the CPS rather than the CEX (recall Section 3).

earnings. Since our model permits an arbitrary correlation between ϵ_k and observables, it will of course just fit the retirement dates of these households. Suppose, however, we replace the residual of each of these households with the mean residual for its quantile and then simulate their retirement ages. Comparing these predicted retirement ages with the actual ages would provide one way of assessing the predictive power of the model's basic mechanisms in the absence of a residual. Figure 3 therefore also graphs the cumulative distribution of retirement ages calculated from Proposition 4 for the same couples after replacing each ϵ_k with the appropriate mean.

Judging by Figure 3, the basic mechanisms of the model appear to simulate retirement behavior reasonably well, with some notable exceptions. For the lowest two quartiles of the income distribution, the model's essential elements predict a bit too much retirement at around ages 56-58 and too little around ages 63-65. The model's basic mechanisms capture the retirement behavior of those in the third quartile of earnings quite well, with the exception of overshooting at age 59. For the fourth quartile of earnings, the model's basic features do very well before age 60, but then predict too little retirement in the mid-60s.

[**Fig 3 here**]

In summary, despite the simplicity of our model, the baseline estimates fit the within-sample data reasonably well. In addition, although aggregative shocks may be important to both expenditure and retirement behavior, our various attempts to accommodate such shocks do not appreciably affect estimates of the parameters of most interest in this paper.

5. Social Security Reform

In this section, we use the estimated model to investigate the consequences of a Social Security reform in which the OASI tax, and all OASI benefit adjustments based on new earnings, cease at a specific age or following a specific span of career years.[26] Although individuals may retire at any age, we assume that those who continue working after the Social Security vesting age/span enjoy a 10.6 percent increase in their aftertax wage.[27] We treat wages and interest rates as exogenous to the model and, thus, unaffected by the reform.

On the benefits side of Social Security, we assume the year-2000 benefits formula.[28] On the tax side, we require that the reform be revenue neutral in the following sense: although lifetime Social Security taxes, earning caps, and indexing remain as in the historical record,

[26] In our model retirement is irreversible and the work day is indivisible. We thus abstract from concerns about how much work is required to be credited with a year towards vesting. In a more elaborate model, such issues would require careful consideration since they would generate incentives for labor supply.

[27] Thus we implicitly assume competitive labor markets in which the incidence of the payroll tax falls entirely on workers.

[28] We assume, however, that the reform decreases the averaging period for the AIME to exclude years past the vesting period.

we calculate an incremental tax rate that, if applied uniformly at all ages, would preserve the historical difference between lifetime OASI benefits and Social Security taxes, computed as present value at household age 50. We impose this extra proportional tax, denoted τ^c, on earnings (up to the cap) prior to the vesting age. (Table 6, below, displays the levels of this extra tax.) Preserving the historical difference between benefits and taxes means preserving the weighted average of benefits net of taxes over the households in a given simulation.

In our framework, a transition to reform could be straightforward. Government could announce that henceforth all new labor–force participants would be subject to the new system — with existing workers remaining in the old system. The revenue neutrality of the proposed reform means that no special complications would arise from "legacy costs" or other intergenerational transfers.

The consequences of the reform for a given household will depend on, among other things, the household's lifetime earning profile and the path of its family structure. Our simulations need to take this variety into account. To preserve realistic heterogeneity, we use our HRS sample of couples. We ask: If each couple had, from birth, faced a reformed Social Security system, how, according to our model, would its retirement age have changed? In other words, we ask what the behavioral responses of the HRS cohorts would have been had government instituted this paper's reform as they began their careers.

The HRS has 3 types of households (recall Table 2): (i) households that voluntarily retire in sample; (ii) households that stop work due to disability; and, (iii) households that work continuously within the sample, never reaching retirement. Our simulations include types (i)-(ii) but drop households of type (iii). We cannot estimate an ϵ_k for the last group.[29]

Our estimation strategy did not require us to specify out-of-sample earnings; this is necessary, however, for purposes of simulation. The simulation results reported in Table 5 rely on the same estimated earnings dynamics equations as in Section 4, where we specified quartic experience profiles to impute top-coded or missing earnings for the HRS. Here we justify fitted values to the last observable HRS earnings figure and then impute earnings to all subsequent ages. (Table 5D — see below — examines the sensitivity of results to this specification.)

Our simulations work as follows. Consider a type (i) household k. Using data on R_k, compute $\epsilon_k = g_k^2$, as in Section 4. Then change the policy regime, which changes the components of g_k^2, giving us, say, g_k^{2*}. We find R_k^* making $g_k^{2*} = \epsilon_k$. A potential complication arises if the vesting age after reform is greater than R_k^*. Then we check for a second critical point, say, R_k^{**} greater than the new vesting age. If a second critical point exists, we choose between R_k^* and R_k^{**} on the basis of the one yielding the highest lifetime utility. The same steps apply for type (ii) households, with q_k^{2*} replacing ϵ_k.

Table 5 presents average simulation outcomes for different vesting ages/spans, and for the two different definitions of disability. These simulations use parameter values from the left-hand side of Table 4.[30] For each row of Table 5, we present a simulation based on

[29] Table 4C, right-hand side, shows that dropping type (iii) households has little impact on our parameter estimates.

[30] The simulation results are qualitatively quite similar if we use point estimates gener-

Table 4's point estimates. Then we draw 1000 random samples for $\widehat{\beta}$, using the asymptotic covariance matrix of the estimates. The table reports medians and 95 percent confidence intervals.

The first column of Table 5 gives the average change in career length as a result of the reform. The second column presents the average equivalent variation of the reform. It shows what, on average, workers in the model would be willing to pay (measured in present value at age 50) in order live their whole lives under the reformed system. This is not, however, the entirety of the efficiency gain from the reform. One must also include additional *income* tax revenue generated by the longer average work careers. This additional income tax revenue, again measured in net present value at age-50, is presented in column 3.

Table 5 reveals three important findings. First, at younger vesting ages, or shorter careers spans, the simulated reform produces economically substantial efficiency gains from large average changes in career length. These gains and changes fall off sharply, however, as vesting ages or career spans rise. At a vesting age of 54 (career span of 34 years) for example, the model predicts an increase in average career length of 1.6-1.7 years, and average equivalent variations of \$4,000-\$4,300. At a vesting age of 64 (career span of 44), however, there is virtually no change in average career length and average equivalent variations diminish to \$100-1000.

A second important finding in Table 5 is that the income tax revenue gains from the reform tend to be substantially larger than the equivalent variations. At a vesting age of 54 (span of 34) the average addition to income tax revenue is \$14,800-\$15,700, more than three times the magnitude of the equivalent variation. In this sense, the proposed reform appears to work best for government solvency, with more modest gains accruing directly to households.

The third key finding revealed in Table 5 is that the fraction of the population that is disabled is an important determinant of the reform's net effects. The top two panels of Table 5 adopt our "stringent," and conventional, definition of disability. The bottom two panels repeat the analysis for our "broad definition." The latter re-classifies many retirements as involuntary. In our framework, households that retire involuntarily cannot respond to new incentives for longer careers. If the broad definition were judged as more accurate, Table 5 shows that gains in labor supply from reform would tend to be about half as great at each vesting age or career span.

A nice feature of our model is that its analytic tractability allows us to understand, with unusual clarity, the mechanisms behind the simulated behavioral responses. To illustrate, focus on the results for the reform with a vesting age of 54. The simulations indicate an average increase in career length of 1.69 years. To see where this response comes from, recall the first-order condition governing optimal retirement (15), rewritten below

$$\frac{B_i'(R) \cdot e^{r \cdot R} + P(R) \cdot y_{R-}}{x_{iR-}} = \frac{1}{\gamma} \cdot P(R) \cdot (\Lambda - 1) \cdot (1 - \gamma) \,. \tag{15'}$$

The left-hand side of (15') gives the benefit of continued work, while the right-hand side gives the utility gain from immediate retirement. Absent the reform, the average

ated by alternative specifications. See below.

retirement age is about 60. So, if the vesting age is 54, a large majority of the simulated households experience a 10.6% increase in last earnings (y_{R-}). Since last earnings are, on average, approximately $70,000 (in 2005 dollars), this amounts to an increase of about $7,500. Because the reform effectively eliminates the incremental Social Security benefits from additional earnings, $B_i'(R)\cdot e^{r\cdot R}$ declines from about $1,950 to about $250, on average. Absent any changes in behavior, the reform thus causes the numerator on the left-hand side of (15') to increase by about $6,000 on average. Except for changes in the probability of disability $P(R)$, the right-hand side of (15') does not change; it reflects just preference parameters. Thus, adjustments to labor supply and consumption are needed to bring (15') back into balance. Postponing retirement by nearly 1.7 years is sufficient to bring (y_{R-}) down by about $4,500; and the additional 1.7 years of earnings translate into more lifetime consumption. In particular, x_{R-} increases by approximately $2,000.

5.1 Heterogenous Welfare Effects

The preceding description of the average consequences of the reform points to heterogeneity in its welfare effects. The reform raises the OASI payroll tax during the vesting period and eliminates it thereafter. Holding pre–reform behavior fixed, the payroll tax changes thus induce what we might call "transfers." Because households that retire earlier than average will not benefit sufficiently from lower taxes late in life to offset their higher payroll tax in youth, their transfers are negative. Conversely, households that retire later than average receive positive transfers. Because reform is revenue neutral, the average transfer is 0. But the reform also generates efficiency gains — virtually exclusively for those who retire voluntarily after vesting. They are the reform's big winners.

Table 6 describes some of this heterogeneity in the welfare consequences of the reform. The first two columns display average effects by the pre–reform retirement age of the household; the second pair of columns describes the average effects by the education level of the household head. Underscoring the intuition developed above, we see that households that, pre–reform, would have retired before vesting, change their labor supply relatively little and suffer sometimes substantial average utility losses from the reform. Those that, absent the reform, would have retired after vesting respond with relatively large career extensions and enjoy substantial average utility gains as a result.

The second two columns of Table 6 show that *who* these winners and losers are depends on the details of the reform. If vesting is determined by age, those with less education do relatively poorly. If, on the other hand, vesting is determined by career span, those with less education do relatively well. This result is driven by a simple relationship. Those with less education tend to have relatively long careers but, because they start their careers at younger ages, they also retire younger.

Tables 5B-D in the appendix examine the sensitivity of our simulation results to assumptions. Our baseline model assumes that a couple retires when the male does. The left-hand side of Table 4C re-estimates parameters treating the earnings of female spouses as exogenous, and Table 5C provides corresponding simulations. Given the birth cohorts of our HRS households, the effects are predictably modest.

Table 5D considers a more elaborate specification of out-of-sample earnings. Instead of predicting out-of-sample earnings from our earnings dynamics equations, we fit a separate quadratic function to each male's last 5 annual (log) earnings figures, constraining the quadratic to fit the last observation perfectly and to have a non-positive slope at the last observed earnings observation. Then we predict out-of-sample HRS male earnings with the quadratic's fitted values. Possible advantages are that the latter predictions reflect idiosyncrasies of the individuals, and take detailed account of very-late-in-life earnings. The effects of reform in Table 5D are about one-third smaller than those of Table 5. They remain, nevertheless, substantial.

6. Conclusion

Certain assumptions of our model — such as jobs requiring full-time work, the permanence of retirement, and the absence or insurability of many forms of risk — likely amplify the behavioral consequences and efficiency gains from reform we have studied. Overall, however, we think the simulated responses of Table 5 are large enough to justify more attention to Social Security reform of this type. In an era of greater longevity, it is paradoxical that older workers do not appear to have extended their careers by much. If distortions from tax policy contribute to this outcome, we think Social Security reform of the type analyzed here might provide an attractive remedy. Indeed, our increases in labor supply would be even larger if the entire payroll tax (rather than just its OASI component) were discontinued at our vesting age.

Although this paper's life–cycle framework is simple, its stripped–down nature enables us to estimate parameters rather than relying on calibration. Similarly, our simulations analyze the behavioral responses from individual households in the HRS sample — which can preserve a realistic degree of heterogeneity of earning profile shapes and family composition. The simulations take into account the standard errors on our parameter estimates.

One of the greatest changes in life styles in recent decades has been the growing labor force participation of married women. While our formulation incorporates both male and female labor supply, one of the elaborations that we are most anxious to pursue is to model female labor force participation decisions in much greater detail.

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Appendix I: Proofs and Details of the Model with Disability

Let $p(t)$ be the probability that a household becomes disabled (or dies) at age t. Letting P_s be the probability of becoming disabled after age s, we have

$$P_s \equiv \int_s^T p(t)\, dt\,. \tag{A1}$$

If D_i is age of disability, then

$$\lambda_{is} \equiv \begin{cases} 1\,, & \text{if } s < \min\{D_i\,,\, R_i\}\,, \\ \lambda > 1\,, & \text{if } s \geq \min\{D_i\,,\, R_i\}\,. \end{cases}$$

Hence, if household i is disabled or retired, it gains utility flow, say, $\bar{u}(x\,,\, s)$ from age-s consumption expenditure x, whereas if it is neither disabled nor retired, it gains a different flow, say, $u(x\,,\, s)$. In the context of our model,

$$\bar{u}(x\,,\, s) = \lambda^{\gamma} \cdot u(x\,,\, s)\,. \tag{A2}$$

Suppress the second argument of $u(.,.)$ in this appendix for expositional convenience.

All households are risk averse because $\gamma < 1$. Hence, after a household plans its retirement age R_i, it fully insures its earnings for ages $s \in [S_i\,,\, R_i]$. Because its desired consumption expenditure may change at age $\min\{D_i\,,\, R_i\}$, a household will fully insure its consumption expenditure stream for ages $s \in [S_i\,,\, R_i]$ as well. (Insurance is superfluous for ages $s > R_i$.)

Dropping the subscript i for expository convenience, let $y_s\,,\ s \in [S\,,\, R]$, be a household's earnings, x_s its consumption expenditure for $s < \min\{D\,,\, R\}$, and $\bar{x}_s$ its expenditure thereafter. The insured values of the household's lifetime earnings, $Y(R)$, and consumption expenditures up to retirement, $X(R)$, are

$$\begin{aligned} Y(R) = \int_S^R p(s) \cdot \int_S^s e^{-r \cdot t} \cdot y_t\, dt\, ds + P_R \cdot \int_S^R e^{-r \cdot t} \cdot y_t\, dt = \\ \int_S^R \int_t^R p(s) \cdot e^{-r \cdot t} \cdot y_t\, ds\, dt + P_R \cdot \int_S^R e^{-r \cdot t} \cdot y_t\, dt = \\ \int_S^R P_t \cdot e^{-r \cdot t} \cdot y_t\, dt\,, \end{aligned} \tag{A3}$$

$$\begin{aligned} X(R) = \int_S^R p(s) \cdot [\int_S^s e^{-r \cdot t} \cdot x_t\, dt + \int_s^R e^{-r \cdot t} \cdot \bar{x}_t\, dt]\, ds + P_R \cdot \int_S^R e^{-r \cdot t} \cdot x_t\, dt = \\ \int_S^R [P_t \cdot e^{-r \cdot t} \cdot x_t + (1 - P_t) \cdot e^{-r \cdot t} \cdot \bar{x}_t]\, dt\,. \end{aligned} \tag{A4}$$

Proposition 1 (2) is a special case of Proposition 3 (4) with $P_s = 1$ all s.

Proof of Proposition 3: Using the notation above, our household's lifetime problem is

$$\max_{x_s, \bar{x}_s\, R} \int_S^R [P_s \cdot e^{-\rho \cdot s} \cdot u(x_s) + (1 - P_s) \cdot e^{-\rho \cdot s} \cdot \bar{u}(\bar{x}_s)]\, ds + \varphi(a_R\,,\, R) \tag{A5}$$

$$\text{subject to:} \quad \dot{a}_s = r \cdot a_s + P_s \cdot y_s - P_s \cdot x_s - (1 - P_s) \cdot \bar{x}_s\,, \tag{A6}$$

$$a_S = 0\,,$$

where post-retirement utility is

$$\varphi(A\,,\, R) \equiv \max_{\bar{x}_s} \int_R^T e^{-\rho \cdot s} \cdot \bar{u}(\bar{x}_s)\, ds \tag{A7}$$

$$\text{subject to:} \quad \dot{a}_s = r \cdot a_s - \bar{x}_s\,, \tag{A8}$$

$$a_{R+} = A + B(R) \cdot e^{r \cdot R}\,, \tag{A9}$$

$$a_T \geq 0\,.$$

(Recall that the age-0 present value of the household's social security benefits is $B(R)$.)

Fix any R. Set up a Hamiltonian for the pre-retirement problem:

$$\mathcal{H} \equiv e^{-\rho \cdot s} \cdot [P_s \cdot u(x_s) + (1 - P_s) \cdot \bar{u}(\bar{x}_s)] + \mu_s \cdot [r \cdot a_s + P_s \cdot y_s - P_s \cdot x_s - (1 - P_s) \cdot \bar{x}_s]\,. \tag{A10}$$

The costate variable μ_s satisfies

$$\dot{\mu}_s = -\frac{\partial \mathcal{H}}{\partial a_s} = -r \cdot \mu_s\,, \tag{A11}$$

$$\mu_R = \frac{\partial \varphi(a_{R-}\,,\, R)}{\partial a_{R-}}\,. \tag{A12}$$

From (A11), we have

$$\mu_s = \mu_0 \cdot e^{-r \cdot s}\,. \tag{A13}$$

First-order conditions yield

$$e^{-\rho \cdot s} \cdot P_s \cdot u'(x_s) = P_s \cdot \mu_s\,, \tag{A14}$$

$$e^{-\rho \cdot s} \cdot (1 - P_s) \cdot \bar{u}'(\bar{x}_s) = (1 - P_s) \cdot \mu_s\,. \tag{A15}$$

The costate for the post-retirement problem also satisfies (A12), and necessary conditions then show

$$e^{-\rho\cdot R}\cdot \bar{u}'(\bar{x}_{R+}) = \frac{\partial\varphi(a_{R-}\,,\,R)}{\partial a_{R-}}\,. \tag{A16}$$

Expressions (A13)-(A14) establish (11); (A13) and (A15) — or (A12) and (A15)-(A16) — establish (13). For $D = \min\{D\,,\,R\}$, (A14)-(A15) show $u'(x_s) = \bar{u}'(\bar{x}_s)$, which establishes (12). If $R = \min\{D\,,\,R\}$, (A12), (A14), and (A16) show the same, again establishing (12). ▮

Proof of Proposition 4: We next make the end time for the pre-retirement problem a free variable. Kamien and Schwartz [1981, p.147] show that we then have an additional necessary condition

$$0 = \mathcal{H}|_{s=R} + \frac{\partial\varphi(A\,,\,R)}{\partial R}|_{A=a_{R-}}\,. \tag{A17}$$

Step 1. Let $A \equiv a_{R-}$. The post-retirement problem in the proof of Proposition 3 is a simpler version of the pre-retirement problem. Letting

$$\psi_s \equiv e^{\frac{r-\rho}{1-\gamma}\cdot s}\,,$$

we can see that the optimal post-retirement expenditure path is

$$\bar{x}_s = \bar{x}\cdot\psi_s\,.$$

The post-retirement budget yields

$$\bar{x}\cdot\int_R^T e^{-r\cdot s}\cdot\psi_s\,ds = A\cdot e^{-r\cdot R} + B(R)\,. \tag{A18}$$

Differentiating,

$$\frac{\partial\bar{x}}{\partial A}\cdot\int_R^T e^{-r\cdot s}\cdot\psi_s\,ds = e^{-r\cdot R}\,, \tag{A19}$$

$$\begin{aligned} &-e^{-r\cdot R}\cdot\bar{x}\cdot\psi_R + \frac{\partial\bar{x}}{\partial R}\cdot\int_R^T e^{-r\cdot s}\cdot\psi_s\,ds = -r\cdot A\cdot e^{-r\cdot R} + B'(R) \Longleftrightarrow \\ &\frac{\partial\bar{x}}{\partial R}\cdot\int_R^T e^{-r\cdot s}\cdot\psi_s\,ds = e^{-r\cdot R}\cdot\bar{x}_{R+} - r\cdot e^{-r\cdot R}\cdot A + B'(R)\,. \end{aligned} \tag{A20}$$

The maximized post-retirement criterion is

$$\varphi(A\,,\,R) = [\bar{x}]^{\gamma}\cdot\int_R^T e^{-\rho\cdot s}\cdot u(\psi_s)\,ds\,,$$

where $\bar{x}$ satisfies (A18). Differentiating,

$$\frac{\partial\varphi}{\partial A} = \gamma\cdot[\bar{x}]^{\gamma-1}\cdot\frac{\partial\bar{x}}{\partial A}\cdot\int_R^T e^{-\rho\cdot s}\cdot u(\psi_s)\,ds\,, \tag{A21}$$

$$\frac{\partial \varphi}{\partial R} = \gamma \cdot [\bar{x}]^{\gamma-1} \cdot \frac{\partial \bar{x}}{\partial R} \cdot \int_R^T e^{-\rho \cdot s} \cdot u(\psi_s)\, ds - [\bar{x}]^{\gamma} \cdot e^{-\rho \cdot R} \cdot u(\psi_R) =$$
$$\frac{\frac{\partial \bar{x}}{\partial R} \cdot \frac{\partial \varphi}{\partial A}}{\frac{\partial \bar{x}}{\partial A}} - e^{-\rho \cdot R} \cdot u(\bar{x} \cdot \psi_R) =$$
$$\frac{\partial \varphi}{\partial A} \cdot e^{r \cdot R} \cdot [e^{-r \cdot R} \cdot \bar{x}_{R+} - r \cdot e^{-r \cdot R} \cdot A + B'(R)] - e^{-\rho \cdot R} \cdot u(\bar{x}_{R+}) . \tag{A22}$$

Expressions (A12) and (A15)-(A16) imply $\bar{x}_{R+} = \bar{x}_{R-} \equiv \bar{x}_R$; (A12) and (A14) imply

$$\frac{\partial \varphi}{\partial A} = \mu_R = e^{-\rho \cdot R} \cdot u'(x_{R-}) ;$$

a_s is continuous, so $a_{R-} = a_R$; x_s ceases to exist at R, so we can write $x_{R-} \equiv x_R$; thus, we can rewrite (A22) as

$$\frac{\partial \varphi}{\partial R} = e^{-\rho \cdot R} \cdot u'(x_R) \cdot [\bar{x}_R - r \cdot a_R + e^{r \cdot R} \cdot B'(R)] - e^{-\rho \cdot R} \cdot u(\bar{x}_R) . \tag{A23}$$

Step 2. Substituting (A10) and (A23) into (A17), we have

$$\begin{aligned} 0 = {} & e^{-\rho \cdot R} \cdot [P_R \cdot u(x_R) + (1 - P_R) \cdot \bar{u}(\bar{x}_R)] + \\ & e^{-\rho \cdot R} \cdot u'(x_R) \cdot [r \cdot a_R + P_R \cdot y_R - P_R \cdot x_R - (1 - P_R) \cdot \bar{x}_R] + \\ & e^{-\rho \cdot R} \cdot u'(x_R) \cdot [\bar{x}_R - r \cdot a_R + r^{r \cdot R} \cdot B'(R)] - e^{-\rho \cdot R} \cdot u(\bar{x}_R) . \end{aligned}$$

Canceling identical terms, we have

$$P_R \cdot [\bar{u}(\bar{x}_R) - u(x_R)] = u'(x_R) \cdot [e^{r \cdot R} \cdot B'(R) + P_R \cdot (y_R - x_R + \bar{x}_R)] . \tag{A24}$$

This completes the proof. ▮

Proof of the Corollary to Proposition 4: Let $\Lambda \equiv [\lambda]^{\frac{\gamma}{1-\gamma}}$. Proposition 3 shows $\Lambda \cdot x_s = \bar{x}_s$. Algebra shows $[\Lambda]^{\gamma} \cdot [\lambda]^{\gamma} = \Lambda$. So, using (A2), we have

$$\bar{u}(\bar{x}_R) - u(x_R) = ([\Lambda]^{\gamma} \cdot [\lambda]^{\gamma} - 1) \cdot u(x_R) = (\Lambda - 1) \cdot u(x_R) .$$

Hence, (A24) yields

$$\begin{aligned} & P_R \cdot (\lambda - 1) \cdot u(x_R) = \frac{\gamma \cdot u(x_R)}{x_R} \cdot [e^{r \cdot R} \cdot B'(R) + P_R \cdot (y_R - x_R + \Lambda \cdot x_R)] \iff \\ & \frac{e^{r \cdot R} \cdot B'(R) + P_R \cdot y_R}{x_R} + P_R \cdot (\Lambda - 1) = \frac{1}{\gamma} \cdot P_R \cdot (\Lambda - 1) \iff \\ & \frac{e^{r \cdot R} \cdot B'(R) + P_R \cdot y_R}{x_R} = \frac{1}{\gamma} \cdot P_R \cdot (\Lambda - 1) \cdot (1 - \gamma) , \end{aligned}$$

which establishes the corollary. ▮

Appendix II: CEX Data

We divide the NIPA and CEX data into 11 categories: food, apparel, personal care, shelter, household operation, transportation, medical care, recreation, education, personal business, and miscellaneous. See Laitner and Silverman [2005] for details. Assuming NIPA data is the more accurate, we scale the CEX data by year and category to the corresponding NIPA amount, applying each scaling factor across every age in the CEX.

This paper uses 3 additional adjustments, as follows.

(1) We subdivide "shelter" into "services from own house" and "other." We scale the latter as we do other categories, but we drop the CEX "services from own house" and impute a substitute that allocates the annual NIPA total service flow from residential houses to the CEX in proportion to CEX reported house values.

(2) CEX medical expenditures omit employer contributions to health insurance and services that Medicare covers. We annually, proportionately, and for every age, adjust CEX expenditures on private health insurance to match the Department of Health and Human Services total for all premiums for private health insurance; and, we adjust out–of–pocket health spending from the CEX to match annual DHHS totals.[31] Turning to Medicare, funding for the benefits comes from a hospital insurance (HI) tax on wages and salaries, monthly premiums for supplementary medical insurance (SMI) from people currently eligible for benefits, and contributions from general tax revenues to SMI. The CEX registers only SMI premiums from participants; so, we allocate the yearly total of Medicare benefits (both HI and all SMI expenditure) to the CEX sample in proportion to SMI premium payments (principally for people over 65).[32]

(3) The NIPA "personal business" category includes bank and brokerage fees, many of which are hidden in the form of low interest on saving accounts, etc., and hence absent from expenditures that CEX households perceive. We assume that bank and brokerage fees make their way into the life–cycle model in the form of lower–than–otherwise interest rates on saving; therefore, we normalize annual personal business expenditures measured in the CEX to match the corresponding NIPA amount less bank and brokerage fees.

Appendix III: HRS Data

We derive male lifetime earnings as follows. Some male earnings figures are missing (e.g., non–FICA employment); the data are right censored at the Social Security tax cap prior to 1980; and, they are right censored at $125,000 for earnings 125,000-250,000, at $250,000 for earnings 250,000-500,000, and at $500,000 for earnings 500,000+ for 1981-1991. Thus, for men we estimate a so–called earnings dynamics model of earnings, dividing the total HRS sample into 4 education groups, and regressing log constant–dollar earnings on a quartic in age and dummy variables for time. The regression error has an individual

[31] See http://www.cms.hhs.gov/statistics/burden-of-health-care-costs/table01.asp. The annual figures cover 1987-2000. We extrapolate to 1984-86 and 2001 using the growth rate of NIPA total medical consumption.

[32] For HI expenditures, see *Social Security Bulletin, Annual Statistical Supplement 2001*, table 8.A1; for SMI receipts and receipts from participant premiums, see table 8.A2.

effect as well as a random term. The likelihood function takes censoring into account. Laitner and Silverman [2005] present details. Prior to 1991, we impute censored observations from the regression. To protect against non–FICA earnings, we also impute from the regression when our data differ from it by over two–thirds of a standard error (for the regression).

After 1991, survey data is available biennially. As a protection against coding errors, we exclude survey earnings greater than twice, or less than half, the earnings dynamics equation prediction for the same age, substituting imputations from the earnings dynamics equation in their place.

Although we use similar steps for female earnings, there are several differences. A woman who never works remains in our sample. As stated above, we assume a woman retires when her spouse does (except for Table 4C, left-hand side, where the earnings of female spouses are taken to be exogenous). We extrapolate non–zero late–in–life earnings only for women who supply market hours in the survey in the last year that their husband works. We are much more concerned than for men that women might have part–time earnings. Prior to 1992, therefore, a woman's earnings are her SSA earnings unless the latter are censored, in which case we impute from female earnings dynamics equations resembling the men's (see Laitner and Silverman [2005]). The HRS provides information in 1996 on whether a woman had non–FICA earnings prior to 1992. If a woman had non–FICA jobs and provided beginning and end dates, we impute her earnings from our earnings–dynamics regressions; if she provided only the span of non–FICA employment, we subtract non–FICA employment years 1980-91, which are evident from the data, and probabilistically impute remaining years using our earnings–dynamics regressions; if a woman said she had non–FICA employment but provided no information on when or how long, we drop the couple from the sample on the basis of incomplete data.

Table 4. Estimated Coefficients Equations (19)-(20): Estimated Parameter (Std. Error/T Stat.)

Parameter or Observation Count	Specification of (19):[a, b]			
	No Time Dummies		Time Dummies in Eq. (19)[c]	
	Stringent Def. Male Disability	Broad Def. Male Disability	Stringent Def. Male Disability	Broad Def. Male Disability
Equation (19): Estimates of β^{CEX} from CEX Data				
$\beta_1^{CEX}=\frac{r-\rho}{1-\gamma}$	0.0264*** (0.0008/34.5720)		0.0279*** (0.0010/29.2975)	
$\beta_2^{CEX}=\beta_3=\xi^S$	0.3351*** (0.0523/6.4107)		0.3066*** (0.0505/6.0678)	
$\beta_3^{CEX}=\beta_4=\xi^K$	0.3372*** (0.0181/18.6686)		0.3363*** (0.0172/19.5193)	
$\beta_4^{CEX}=\frac{\gamma}{1-\gamma}\cdot\ln(\lambda)$	-0.0831** (0.0370/-2.2482)		-0.0750** (0.0352/-2.1299)	
Observations	765		765	
Equation (20): Estimates of β given β^{CEX}; HRS Data; ξ^S $(=\beta_3)$ and ξ^K $(=\beta_4)$ as above				
$\beta_1=\rho$	0.0143*** (0.0016/8.9820)	0.0141*** (0.0017/8.1098)	0.0131*** (0.0016/7.9322)	0.0129*** (0.0018/7.2238)
$\beta_2=\gamma$	-0.0797** (0.0370/-2.1537)	-0.0888** (0.0432/-2.0532)	-0.0693** (0.0343/-2.0211)	-0.0772* (0.0403/-1.9167)
$\beta_5=\lambda$	3.0847*** (0.5328/5.7893)	2.7716*** (0.5751/4.8193)	3.1772*** (0.6437/4.9362)	2.8473*** (0.6919/4.1150)
σ_ϵ^2	0.1783 (0.1912/0.9325)	0.1587 (0.1828/0.8681)	0.1879 (0.2347/0.8007)	0.1664 (0.2217/0.7505)
Observations	924	924	924	924

Source: see text. Significant at * 10%, ** 5%, and *** 1% level.

a. Unless otherwise noted, CEX adult (female) ages 25-69 and no time dummies.

b. Regressor 3, top panel: per household number kids ages 0-17, with weight 0.7, and 18-25 if in school or living with parents, with weight 1.0; up to maximum of 2 — see text.

c. Time–dummy coefficients omitted.

Table 5. Policy Simulations using Estimated Parameters of Table 4, LHS: Mean/Median Value and [95% Confidence Interval][a]			
Vesting Age or Span of Years	Average Change Career Years	Average Equivalent Variation (PV Age 50; 2005 NIPA PCE Dollars)	Average Additional Income Tax Revenue Per Household (PV Age 50; 2005 NIPA PCE Dollars)
Age	Stringent Definition Disability (see Table 2); Vesting by Age		
54	1.69/1.71 [1.61 , 1.82]	4292.82/4197.75 [3886.82 , 4566.05]	15684.74/15652.77 [14447.37 , 16979.26]
59	1.08/1.08 [0.99 , 1.15]	2677.61/2503.94 [1947.75 , 3078.05]	9197.92/9012.78 [8388.53 , 9503.62]
64	-0.02/-0.03 [-0.09 , 0.06]	233.70/89.89 [-513.59 , 701.00]	1385.58/1317.73 [759.79 , 2090.91]
Span	Stringent Definition Disability (see Table 2); Vesting by Career Span		
34	1.63/1.67 [1.57 , 1.79]	4079.26/4066.01 [3753.12 , 4401.39]	14828.08/14933.84 [13910.95 , 16415.43]
39	1.24/1.27 [1.17 , 1.36]	3723.44/3616.23 [2976.30 , 4089.34]	10258.31/10397.80 [9517.31 , 11126.99]
44	0.31/0.32 [0.22 , 0.42]	1024.83/1033.39 [206.04 , 1607.89]	2870.73/2972.30 [2107.92 , 3581.56]
Age	Broad Definition Disability (see Table 2); Vesting by Age		
54	0.94/0.91 [0.75 , 0.99]	2269.83/2159.03 [1715.45 , 2495.10]	8188.89/7972.27 [6340.89 , 8675.11]
59	0.61/0.58 [0.45 , 0.67]	1476.43/1416.73 [1029.65 , 2033.86]	4623.10/4443.38 [3428.68 , 5301.52]
64	-0.11/-0.10 [-0.24 , -0.03]	-873.87/-765.09 [-1402.58 , -138.91]	-117.33/-69.85 [-883.42 , 553.96]
Span	Broad Definition Disability (see Table 2); Vesting by Career Span		
34	0.93/0.89 [0.74 , 0.97]	2192.09/2117.91 [1683.04 , 2379.93]	7786.25/7626.82 [6095.13 , 8161.78]
39	0.62/0.62 [0.49 , 0.74]	1586.06/1620.96 [1209.47 , 2531.16]	4682.04/4718.48 [3655.60 , 5816.34]
44	-0.05/-0.05 [-0.08 , -0.04]	-532.30/-549.03 [-703.61 , -301.08]	-3.90/-7.98 [-234.01 , 146.66]

Source: see text.

a. “Mean” based upon point estimates; median and confidence intervals based on 1000 random parameter vector draws — see text.

Table 6. Additional, Detailed Simulation Output,

Estimated Parameters from Table 4, Column 1

Vesting Age/ Span	Pre-Vesting Payroll Tax Change	**Ave Change Career Yrs/Ave Equiv Variation**[a] **(Fraction of Sample)**			
		Pre-Reform Retirement Age		Education	
		Ret Age ≤ Vesting Age	Ret Age > Vesting Age	High School or less	More Than High School
Age	Stringent Definition Disability (see Table 2); Vesting by Age				
54	0.0093	0.21/-15839.88 (0.10)	1.86/6632.65 (0.90)	1.84/2189.96 (0.47)	1.56/6173.36 (0.53)
59	0.0027	0.11/-2389.04 (0.42)	1.79/6419.92 (0.58)	1.02/429.92 (0.47)	1.13/4676.47 (0.53)
64	0.0001	-0.13/-196.58 (0.92)	1.30/5302.21 (0.08)	-0.41/-1794.65 (0.46)	0.32/1942.74 (0.54)
Span	Stringent Definition Disability (see Table 2); Vesting by Career Span				
34	0.0096	0.86/-13475.36 (0.11)	1.73/6341.34 (0.89)	1.90/13826.98 (0.47)	1.39/-4739.48 (0.53)
39	0.0027	0.79/-779.57 (0.43)	1.58/7067.43 (0.57)	1.51/9332.77 (0.47)	1.00/-1206.10 (0.53)
44	0.0001	0.29/830.38 (0.92)	0.56/3148.03 (0.08)	0.40/558.78 (0.46)	0.24/1414.48 (0.54)
Age	Broad Definition Disability (see Table 2); Vesting by Age				
54	0.0092	-0.14/-17996.62 (0.12)	1.09/4920.81 (0.88)	0.92/82.36 (0.45)	0.97/4048.76 (0.55)
59	0.0026	0.03/-3550.72 (0.43)	1.03/5229.33 (0.57)	0.49/-311.96 (0.45)	0.70/2944.83 (0.55)
64	0.0002	-0.19/-1314.51 (0.92)	0.85/4171.05 (0.08)	-0.28/-1695.17 (0.45)	0.03/-194.42 (0.55)
Span	Broad Definition Disability (see Table 2); Vesting by Career Span				
34	0.0094	0.31/-16309.75 (0.12)	1.01/4737.23 (0.88)	0.99/12249.91 (0.45)	0.87/-6084.58 (0.55)
39	0.0027	0.34/-2969.51 (0.43)	0.84/4980.17 (0.57)	0.71/7803.59 (0.45)	0.55/-3438.53 (0.55)
44	0.0002	-0.03/-870.60 (0.91)	0.30/2483.78 (0.09)	0.03/-559.35 (0.45)	-0.02/-582.92 (0.55)

Source: See text.

a. Present value age 50, 2005 NIPA PCE Dollars.

Figure 1: Average Annual Changes in the CEX, by Age

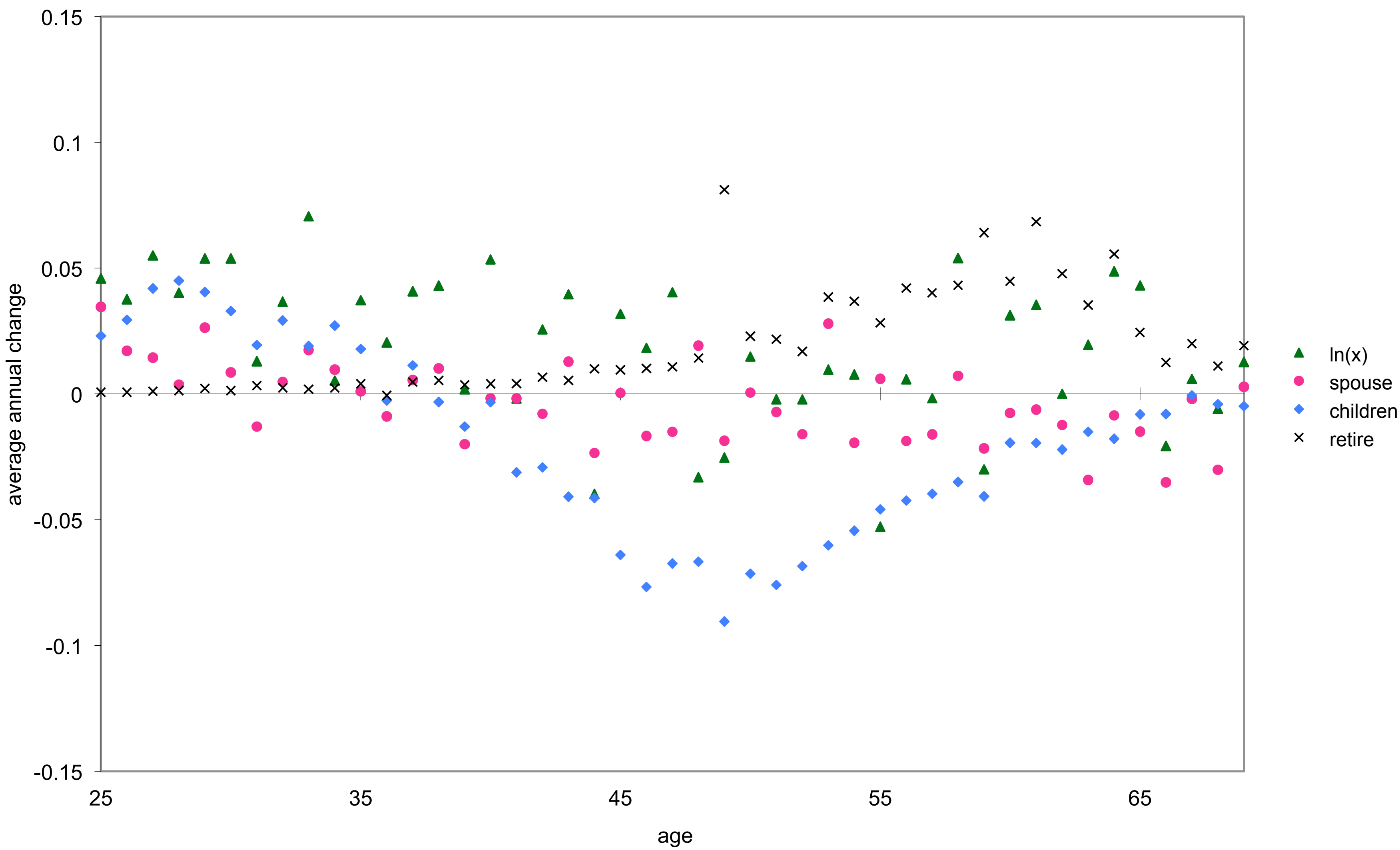

Figure 2: CEX Average Annual Expenditure Change Residuals, by Age

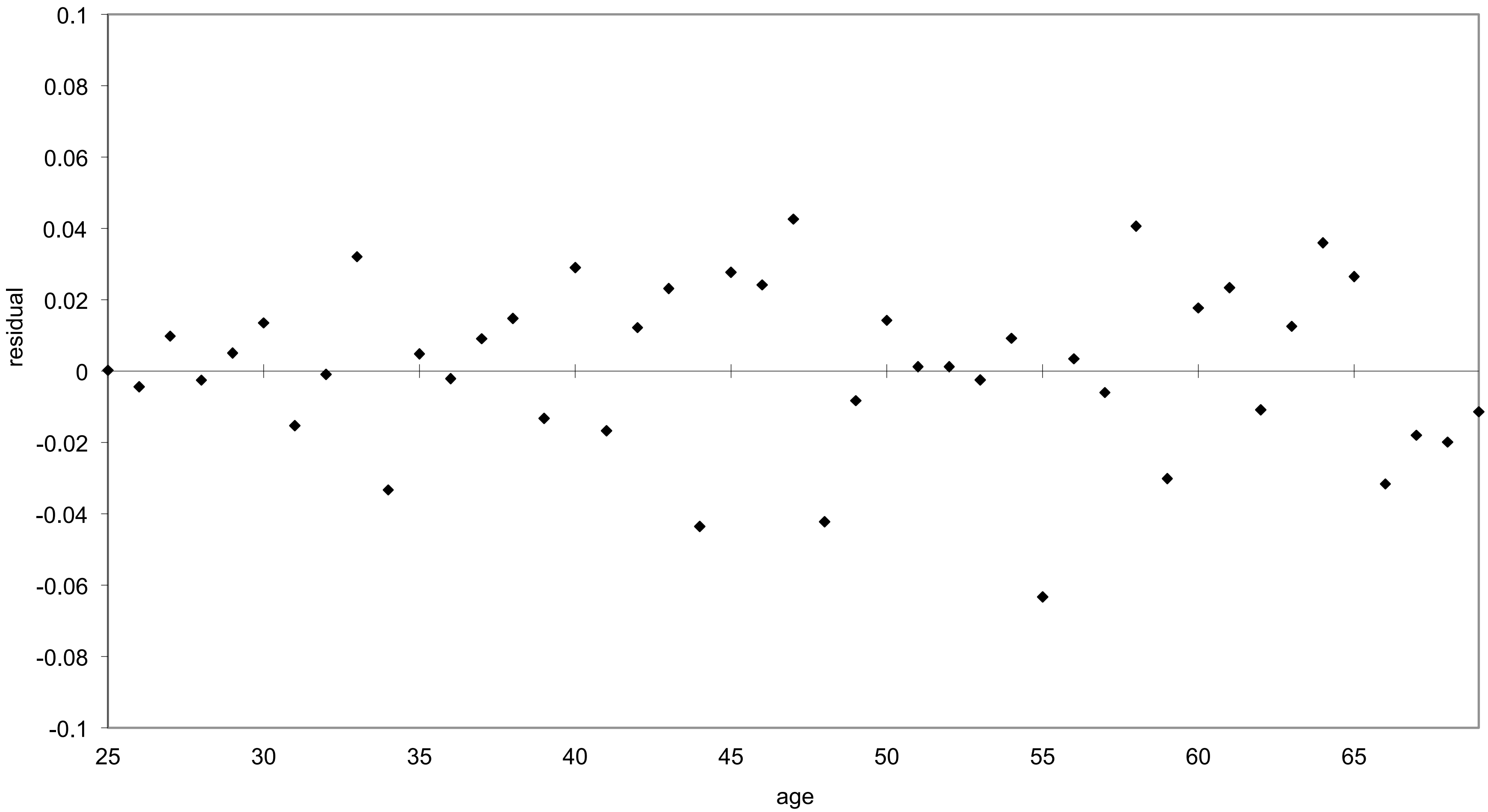

Figure 3a: Cumulative Fraction Retired, HRS 1992-2002 and Simulated Model
First Quartile of Earnings Distribution

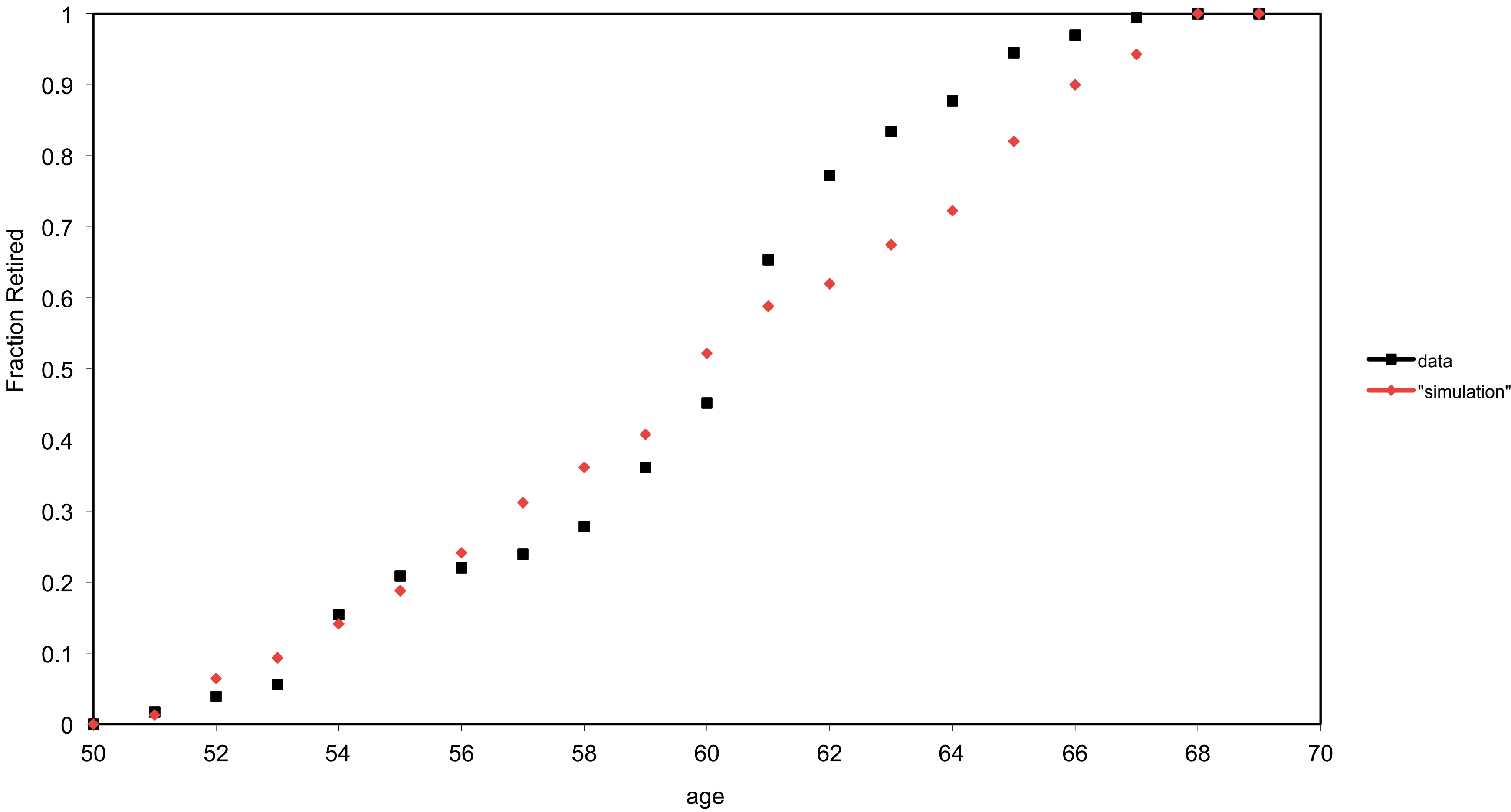

Figure 3b: Cumulative Fraction Retired, HRS 1992-2002 and Simulated Model
Second Quartile of Earnings Distribution

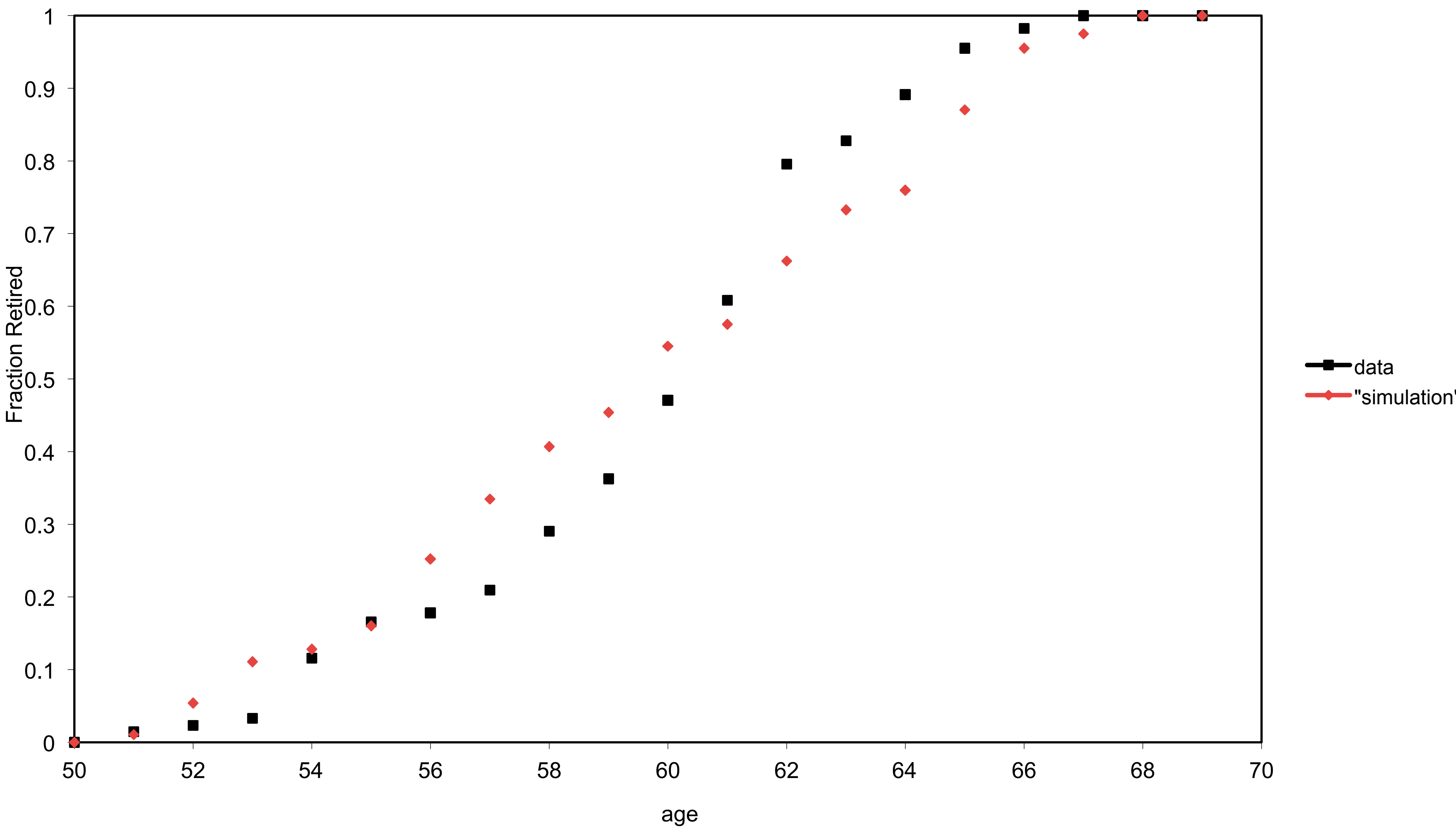

Figure 3c: Cumulative Fraction Retired, HRS 1992-2002 and Simulated Model
Third Quartile of Earnings Distribution

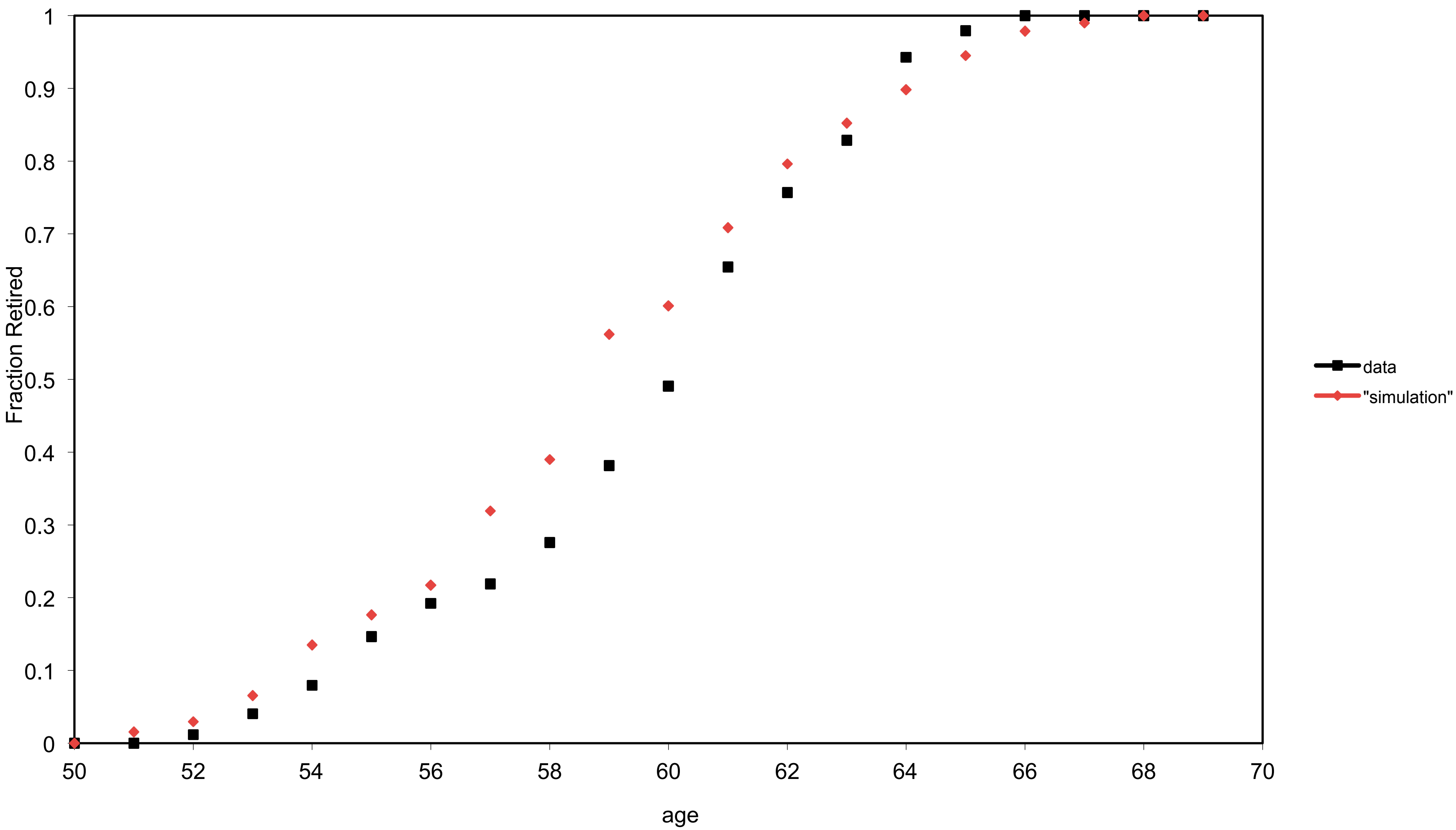

Figure 3d: Cumulative Fraction Retired, HRS 1992-2002 and Simulated Model
Fourth Quartile of Earnings Distribution

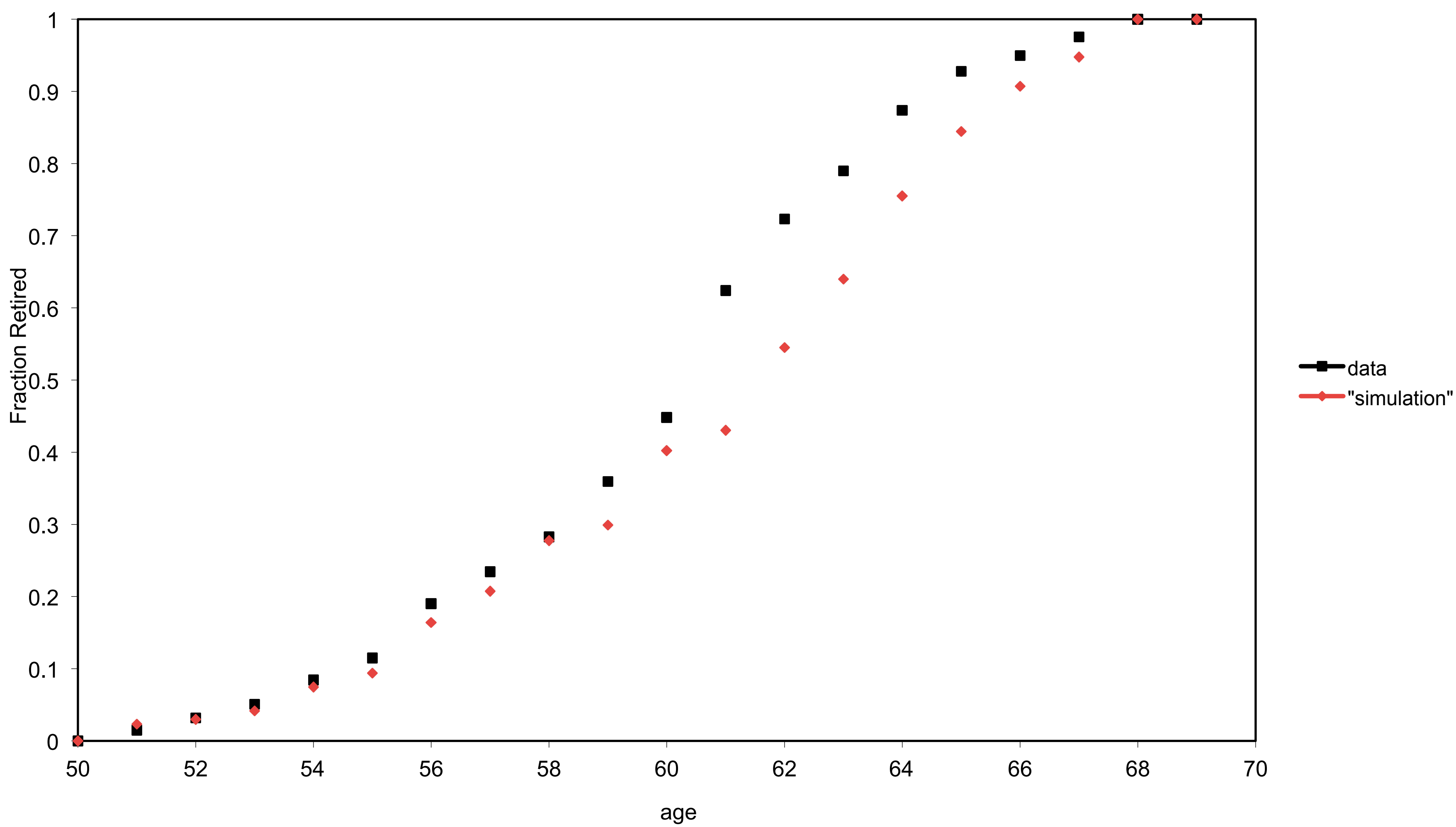

Table 4B. Estimated Coefficients Equations (19)-(20): Estimated Parameter (Std. Error/T Stat.)

Parameter or Observation Count	Specification of (19):[a, b] Time Dummies in Eqs. (19)&(20)[c] Stringent Def. Male Disability	Time Dummies in Eqs. (19)&(20)[c] Broad Def. Male Disability	Aggregation 1984-2000; No Time Dummies Stringent Def. Male Disability	Aggregation 1984-2000; No Time Dummies Broad Def. Male Disability
Equation (19): Estimates of β^{CEX} from CEX Data				
$\beta_1^{CEX}=\frac{r-\rho}{1-\gamma}$	0.0279*** (0.0010/29.2975)		0.0265*** (0.0047/5.6207)	
$\beta_2^{CEX}=\beta_3=\xi^S$	0.3066*** (0.0505/6.0678)		0.5262** (0.2306/2.2818)	
$\beta_3^{CEX}=\beta_4=\xi^K$	0.3363*** (0.0172/19.5193)		0.3243*** (0.1102/2.9429)	
$\beta_4^{CEX}=\frac{\gamma}{1-\gamma}\cdot\ln(\lambda)$	-0.0750** (0.0352/-2.1299)		-0.0023 (0.1940/-0.0117)	
Observations	765		45	
Equation (20): Estimates of β given β^{CEX}; HRS Data; ξ^S $(=\beta_3)$ and ξ^K $(=\beta_4)$ as above				
$\beta_1=\rho$	0.0132*** (0.0016/8.1624)	0.0130*** (0.0017/7.5437)	0.0163* (0.0087/1.8769)	0.0163* (0.0092/1.7725)
$\beta_2=\gamma$	-0.0652* (0.0334/-1.9524)	-0.0714* (0.0384/-1.8615)	-0.0022 (0.1900/-0.0117)	-0.0025 (0.2112/-0.0117)
$\beta_5=\lambda$	3.4064*** (0.8989/3.7895)	3.0781*** (0.9050/3.4013)	2.7814*** (0.7751/3.5885)	2.5108*** (0.7343/3.4192)
σ_ϵ^2	0.1919 (0.2752/0.6972)	0.1548 (0.2353/0.6579)	0.1616 (0.1738/0.9295)	0.1445 (0.1651/0.8753)
Observations	899	899	924	924

Source: see text. Significant at * 10%, ** 5%, and *** 1% level.

a. Unless otherwise noted, CEX adult (female) ages 25-69 and no time dummies.

b. Regressor 3, top panel: per household number kids ages 0-17, with weight 0.7, and 18-25 if in school or living with parents, with weight 1.0; up to maximum of 2 — see text.

c. Annual time-dummies for eq. (19); HRS birthyear-dummies, 1928-40, for eq. (20). All time–dummy coefficients omitted.

Table 4C. Estimated Coefficients Equations (19)-(20): Estimated Parameter (Std. Error/T Stat.)

Parameter or Observation Count	Specification of (19):[a, b]			
	Female Earnings Treated as Exogenous		Exclude Households of Males Not Retiring in Sample	
	Stringent Def. Male Disability	Broad Def. Male Disability	Stringent Def. Male Disability	Broad Def. Male Disability
Equation (19): Estimates of β^{CEX} from CEX Data				
$\beta_1^{CEX}=\frac{r-\rho}{1-\gamma}$	0.0264*** (0.0008/34.5720)		0.0264*** (0.0008/34.5720)	
$\beta_2^{CEX}=\beta_3=\xi^S$	0.3351*** (0.0523/6.4107)		0.3351*** (0.0523/6.4107)	
$\beta_3^{CEX}=\beta_4=\xi^K$	0.3372*** (0.0181/18.6686)		0.3372*** (0.0181/18.6686)	
$\beta_4^{CEX}=\frac{\gamma}{1-\gamma}\cdot\ln(\lambda)$	-0.0831** (0.0370/-2.2482)		-0.0831** (0.0370/-2.2482)	
Observations	765		765	
Equation (20): Estimates of β given β^{CEX}; HRS Data; ξ^S $(=\beta_3)$ and ξ^K $(=\beta_4)$ as above				
$\beta_1=\rho$	0.0136*** (0.0020/6.9110)	0.0132*** (0.0023/5.8484)	0.0148*** (0.0014/10.8586)	0.0147*** (0.0014/10.1393)
$\beta_2=\gamma$	-0.1082** (0.0517/-2.0947)	-0.1234* (0.0640/-1.9283)	-0.0603** (0.0277/-2.1791)	-0.0665** (0.0312/-2.1301)
$\beta_5=\lambda$	2.3421*** (0.3472/6.7450)	2.1309*** (0.4009/5.3155)	4.3105*** (0.8979/4.8006)	3.7947*** (0.8450/4.4909)
σ_ϵ^2	0.1462 (0.1798/0.8132)	0.1258 (0.1832/0.6870)	0.3478 (0.2785/1.2491)	0.2311 (0.2194/1.0533)
Observations	924	924	444	332

Source: see text. Significant at * 10%, ** 5%, and *** 1% level.

a. Unless otherwise noted, CEX adult (female) ages 25-69 and no time dummies.

b. Regressor 3, top panel: per household number kids ages 0-17, with weight 0.7, and 18-25 if in school or living with parents, with weight 1.0; up to maximum of 2 — see text.

Table 4D. Estimated Coefficients Equations (19)-(20): Estimated Parameter (Std. Error/T Stat.)

Parameter or Observation Count	Specification of (19):[a, b]			
	CEX Adult (Female) Ages 30-69		CEX Adult (Female) Ages 35-69	
	Stringent Def. Male Disability	Broad Def. Male Disability	Stringent Def. Male Disability	Broad Def. Male Disability
Equation (19): Estimates of β^{CEX} from CEX Data				
$\beta_1^{CEX}=\frac{r-\rho}{1-\gamma}$	0.0276*** (0.0010/27.8506)		0.0317*** (0.0016/19.5793)	
$\beta_2^{CEX}=\beta_3=\xi^S$	0.3446*** (0.0591/5.8339)		0.3940*** (0.0640/6.1534)	
$\beta_3^{CEX}=\beta_4=\xi^K$	0.3591*** (0.0220/16.3054)		0.4165*** (0.0285/14.6055)	
$\beta_4^{CEX}=\frac{\gamma}{1-\gamma}\cdot\ln(\lambda)$	-0.0876** (0.0382/-2.2957)		-0.1235*** (0.0408/-3.0241)	
Observations	680		595	
Equation (20): Estimates of β given β^{CEX}; HRS Data; ξ^S $(=\beta_3)$ and ξ^K $(=\beta_4)$ as above				
$\beta_1=\rho$	0.0130*** (0.0018/7.4500)	0.0128*** (0.0019/6.7680)	0.0079*** (0.0024/3.2649)	0.0075*** (0.0026/2.9265)
$\beta_2=\gamma$	-0.0814** (0.0369/-2.2051)	-0.0906** (0.0432/-2.1000)	-0.1028*** (0.0353/-2.9142)	-0.1141*** (0.0424/-2.6878)
$\beta_5=\lambda$	3.2033*** (0.5728/5.5925)	2.8707*** (0.6138/4.6766)	3.7624*** (0.7868/4.7820)	3.3397*** (0.8238/4.0538)
σ_ϵ^2	0.1880 (0.2036/0.9236)	0.1665 (0.1932/0.8617)	0.2282 (0.2615/0.8728)	0.1985 (0.2448/0.8110)
Observations	924	924	924	924

Source: see text. Significant at * 10%, ** 5%, and *** 1% level.

a. Unless otherwise noted, CEX adult (female) ages 25-69 and no time dummies.

b. Regressor 3, top panel: per household number kids ages 0-17, with weight 0.7, and 18-25 if in school or living with parents, with weight 1.0; up to maximum of 2 — see text.

Table 4E. Estimated Coefficients Equations (19)-(20):
Estimated Parameter (Std. Error/T Stat.)

Parameter or Observation Count	Specification of (19):[a, b]			
	CEX Adult (Female) Ages 40-69		CEX Adult (Female) Ages 25-64	
	Stringent Def. Male Disability	Broad Def. Male Disability	Stringent Def. Male Disability	Broad Def. Male Disability
Equation (19): Estimates of β^{CEX} from CEX Data				
$\beta_1^{CEX}=\frac{r-\rho}{1-\gamma}$	0.0325*** (0.0030/10.7826)		0.0268*** (0.0008/35.4742)	
$\beta_2^{CEX}=\beta_3=\xi^S$	0.3901*** (0.0715/5.4570)		0.3389*** (0.0597/5.6738)	
$\beta_3^{CEX}=\beta_4=\xi^K$	0.4283*** (0.0375/11.4075)		0.3164*** (0.0197/16.0514)	
$\beta_4^{CEX}=\frac{\gamma}{1-\gamma}\cdot\ln(\lambda)$	-0.1337** (0.0574/-2.3287)		-0.1523*** (0.0398/-3.8232)	
Observations	510		680	
Equation (20): Estimates of β given β^{CEX}; HRS Data; ξ^S $(=\beta_3)$ and ξ^K $(=\beta_4)$ as above				
$\beta_1=\rho$	0.0068 (0.0043/1.5765)	0.0064 (0.0045/1.4261)	0.0121*** (0.0017/6.9966)	0.0117*** (0.0020/5.9453)
$\beta_2=\gamma$	-0.1084** (0.0429/-2.5296)	-0.1203** (0.0507/-2.3715)	-0.1459*** (0.0436/-3.3480)	-0.1626*** (0.0538/-3.0214)
$\beta_5=\lambda$	3.9232*** (0.9916/3.9566)	3.4736*** (0.9963/3.4866)	3.3097*** (0.5821/5.6854)	2.9724*** (0.6185/4.8059)
σ_ϵ^2	0.2396 (0.2810/0.8524)	0.207 (0.2658/0.7809)	0.1839 (0.2012/0.9141)	0.1633 (0.1909/0.8552)
Observations	924	924	924	924

Source: see text. Significant at * 10%, ** 5%, and *** 1% level.

a. Unless otherwise noted, CEX adult (female) ages 25-69 and no time dummies.

b. Regressor 3, top panel: per household number kids ages 0-17, with weight 0.7, and 18-25 if in school or living with parents, with weight 1.0; up to maximum of 2 — see text.

Table 4F. Estimated Coefficients Equations (19)-(20): Estimated Parameter (Std. Error/T Stat.)

Parameter or Observation Count	Specification of (19):[a, b]			
	CEX Adult (Female) Ages 25-74		CEX Child Ages 0-17 (Weight 0.7) [Exclude Children Ages 18+]; CEX Adult (Female) Ages 25-79	
	Stringent Def. Male Disability	Broad Def. Male Disability	Stringent Def. Male Disability	Broad Def. Male Disability
Equation (19): Estimates of β^{CEX} from CEX Data				
$\beta_1^{CEX}=\frac{r-\rho}{1-\gamma}$	0.0252*** (0.0007/33.7981)		0.0269*** (0.0009/30.8761)	
$\beta_2^{CEX}=\beta_3=\xi^S$	0.4201*** (0.0461/9.1040)		0.4989*** (0.0464/10.7472)	
$\beta_3^{CEX}=\beta_4=\xi^K$	0.3233*** (0.0180/17.9635)		0.2016*** (0.0227/8.8737)	
$\beta_4^{CEX}=\frac{\gamma}{1-\gamma}\cdot\ln(\lambda)$	-0.0542 (0.0368/-1.4729)		-0.3557*** (0.0339/-10.4994)	
Observations	850		935	
Equation (20): Estimates of β given β^{CEX}; HRS Data; ξ^S $(=\beta_3)$ and ξ^K $(=\beta_4)$ as above				
$\beta_1=\rho$	0.0163*** (0.0016/10.5149)	0.0162*** (0.0017/9.6733)	0.0064*** (0.0022/2.9498)	0.0054* (0.0028/1.8936)
$\beta_2=\gamma$	-0.054 (0.0376/-1.4414)	-0.0604 (0.0428/-1.4101)	-0.3525*** (0.0613/-5.7460)	-0.3917*** (0.0901/-4.3477)
$\beta_5=\lambda$	2.8738*** (0.4720/6.0880)	2.5928*** (0.5094/5.0900)	3.9156*** (0.6785/5.7711)	3.5388*** (0.7399/4.7825)
σ_ϵ^2	0.1633 (0.1718/0.9503)	0.1463 (0.1643/0.8904)	0.1860 (0.2152/0.8645)	0.1659 (0.2205/0.7522)
Observations	924	924	924	924

Source: see text. Significant at * 10%, ** 5%, and *** 1% level.

a. Unless otherwise noted, CEX adult (female) ages 25-69; CEX child ages 0-17, with weight 0.7, and 18-25 if in school or living with parents, with weight 1.0; and, children up to maximum of 2 counted.

b. No time dummies unless explicitly indicated.

Table 4G. Estimated Coefficients Equations (19)-(20):
Estimated Parameter (Std. Error/T Stat.)

Parameter or Observation Count	Specification of (19):[a, b]			
	CEX Child Ages 0-17 (Weight 0.7) [Exclude Children Ages 18+]; CEX Adult (Female) Ages 25-69		Children Counted to Maximum of 3	
	Stringent Def. Male Disability	Broad Def. Male Disability	Stringent Def. Male Disability	Broad Def. Male Disability
Equation (19): Estimates of β^{CEX} from CEX Data				
$\beta_1^{CEX}=\frac{r-\rho}{1-\gamma}$	0.0312*** (0.0010/31.7779)		0.0269*** (0.0008/34.8589)	
$\beta_2^{CEX}=\beta_3=\xi^S$	0.2720*** (0.0599/4.5448)		0.3167*** (0.0528/5.9932)	
$\beta_3^{CEX}=\beta_4=\xi^K$	0.2617*** (0.0229/11.4336)		0.2712*** (0.0149/18.1643)	
$\beta_4^{CEX}=\frac{\gamma}{1-\gamma}\cdot\ln(\lambda)$	-0.4989*** (0.0369/-13.5325)		-0.1405*** (0.0360/-3.9022)	
Observations	765		765	
Equation (20): Estimates of β given β^{CEX}; HRS Data; ξ^S $(=\beta_3)$ and ξ^K $(=\beta_4)$ as above				
$\beta_1=\rho$	-0.0005 (0.0026/-0.1966)	-0.0018 (0.0037/-0.4760)	0.0122*** (0.0017/7.2839)	0.0118*** (0.0019/6.1708)
$\beta_2=\gamma$	-0.3899*** (0.0700/-5.5713)	-0.4296*** (0.1086/-3.9563)	-0.1362*** (0.0396/-3.4373)	-0.1518*** (0.0505/-3.0057)
$\beta_5=\lambda$	5.9210*** (1.4545/4.0710)	5.2595*** (1.6076/3.2717)	3.2302*** (0.5587/5.7815)	2.9053*** (0.6294/4.6163)
σ_ϵ^2	0.2717 (0.3907/0.6954)	0.2374 (0.4076/0.5825)	0.1785 (0.1949/0.9160)	0.1588 (0.1970/0.8059)
Observations	924	924	924	924

Source: see text. Significant at * 10%, ** 5%, and *** 1% level.

a. Unless otherwise noted, CEX adult (female) ages 25-69; CEX child ages 0-17, with weight 0.7, and 18-25 if in school or living with parents, with weight 1.0; and, children up to maximum of 2 counted.

b. No time dummies unless explicitly indicated.

Table 5B. Policy Simulations using Estimated Parameters of Table 4B, RHS: Mean/Median Value and [95% Confidence Interval][a]			
Vesting Age or Span of Years	Average Change Career Years	Average Equivalent Variation (PV Age 50; 2005 NIPA PCE Dollars)	Average Additional Income Tax Revenue Per Household (PV Age 50; 2005 NIPA PCE Dollars)
Age	Stringent Definition Disability (see Table 2); Vesting by Age		
54	1.65/1.83 [1.43 , 2.61]	3842.19/4291.89 [2503.89 , 6241.63]	14835.93/16767.15 [12566.91 , 24642.90]
59	1.05/1.09 [0.81 , 1.48]	2174.85/1909.88 [374.28 , 3106.50]	8621.47/8951.29 [5745.41 , 11893.84]
64	-0.06/-0.06 [-0.29 , 0.07]	-596.99/-510.68 [-2607.90 , 1063.37]	703.53/934.24 [-2070.00 , 2340.76]
Age	Broad Definition Disability (see Table 2); Vesting by Age		
54	0.88/0.96 [0.71 , 1.30]	1852.83/2138.89 [944.76 , 4014.50]	7419.65/8308.81 [5764.14 , 11362.99]
59	0.55/0.58 [0.39 , 0.83]	1052.95/1172.70 [-11.98 , 3079.09]	3986.03/4225.53 [2325.81 , 6331.00]
64	-0.13/-0.14 [-0.30 , 0.02]	-1076.30/-1069.84 [-2177.80 , 721.60]	-320.23/-463.85 [-1963.09 , 1241.76]

Source: see text.

a. "Mean" based upon point estimates; median and confidence intervals based on 1000 random parameter vector draws — see text.

Table 5C. Policy Simulations using Estimated Parameters from Table 4C, LHS, and Treating Female Earnings as Exogenous: Mean/Median Value and [95% Confidence Interval][a]			
Vesting Age or Span of Years	Average Change Career Years	Average Equivalent Variation (PV Age 50; 2005 NIPA PCE Dollars)	Average Additional Income Tax Revenue Per Household (PV Age 50; 2005 NIPA PCE Dollars)
Age	Stringent Definition Disability (see Table 2); Vesting by Age		
54	1.53/1.54 [1.39 , 1.66]	2936.60/2837.74 [2410.60 , 3271.16]	10734.14/10646.57 [9268.05 , 11476.97]
59	0.93/0.95 [0.89 , 1.03]	1113.39/1165.81 [797.89 , 1588.02]	5646.13/5723.44 [5401.56 , 6120.14]
64	0.00/-0.02 [-0.10 , 0.08]	-179.45/-354.32 [-947.85 , 213.31]	1557.41/1398.34 [814.97 , 1941.85]
Age	Broad Definition Disability (see Table 2); Vesting by Age		
54	0.85/0.84 [0.69 , 0.92]	1602.85/1517.31 [1156.36 , 1850.30]	5701.64/5575.48 [4383.14 , 6309.51]
59	0.54/0.53 [0.43 , 0.59]	881.23/824.04 [603.89 , 1111.65]	3099.25/3034.81 [2312.52 , 3496.65]
64	-0.09/-0.10 [-0.17 , -0.04]	-542.86/-604.91 [-966.48 , -256.12]	304.14/213.95 [-189.81 , 624.20]

Source: see text.

a. “Mean” based upon point estimates; median and confidence intervals based on 1000 random parameter vector draws — see text.

Table 5D. Policy Simulations using Estimated Parameters from Table 4, LHS, with Post-Retirement Male Earnings Predicted with a Quadratic (see text): Mean/Median Value and [95% Confidence Interval][a]			
Vesting Age or Span of Years	Average Change Career Years	Average Equivalent Variation (PV Age 50; 2005 NIPA PCE Dollars)	Average Additional Income Tax Revenue Per Household (PV Age 50; 2005 NIPA PCE Dollars)
Age	Stringent Definition Disability (see Table 2); Vesting by Age		
54	1.12/1.13 [1.06 , 1.21]	2588.60/2682.19 [2471.57 , 2928.86]	10254.57/10314.15 [9663.51 , 11066.58]
59	0.65/0.64 [0.59 , 0.68]	1063.43/1078.50 [764.64 , 1466.99]	5265.65/5243.81 [4595.35 , 5485.44]
64	-0.23/-0.22 [-0.25 , -0.18]	-1671.11/-1590.87 [-2001.23 , -1143.17]	-1208.96/-1095.16 [-1468.22 , -724.99]
Age	Broad Definition Disability (see Table 2); Vesting by Age		
54	0.69/0.66 [0.56 , 0.70]	1462.71/1417.06 [1180.25 , 1574.73]	5605.87/5472.66 [4609.81 , 5870.84]
59	0.44/0.41 [0.33 , 0.46]	748.96/750.91 [565.84 , 1145.34]	2917.46/2873.64 [2274.07 , 3440.34]
64	-0.17/-0.17 [-0.28 , -0.11]	-1339.67/-1231.12 [-1993.25 , -805.24]	-839.51/-763.40 [-1666.48 , -406.54]

Source: see text.

a. "Mean" based upon point estimates; median and confidence intervals based on 1000 random parameter vector draws — see text.