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THE EFFECTS OF HEALTH INSURANCE AND SELF-INSURANCE ON RETIREMENT BEHAVIOR

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This paper provides an empirical analysis of the effects of employer-provided health insurance, Medicare, and Social Security on retirement behavior. Using data from the Health and Retirement Study, we estimate a dynamic programming model of retirement that accounts for both saving and uncertain medical expenses. Our results suggest that Medicare is important for understanding retirement behavior, and that uncertainty and saving are both important for understanding the labor supply responses to Medicare. Half the value placed by a typical worker on his employer-provided health insurance is the value of reduced medical expense risk. Raising the Medicare eligibility age from 65 to 67 leads individuals to work an additional 0.074 years over ages 60–69. In comparison, eliminating 2 years worth of Social Security benefits increases years of work by 0.076 years.

KEYWORDS: Retirement behavior, saving, health insurance, Medicare.

1. INTRODUCTION

ONE OF THE LARGEST SOCIAL PROGRAMS for the rapidly growing elderly population is Medicare. In 2009, Medicare had 46.3 million beneficiaries and $509 billion of expenditures, making it comparable to Social Security.[2]

Prior to receiving Medicare at age 65, many individuals receive health insurance only if they continue to work. This work incentive disappears at age 65, when Medicare provides health insurance to almost everyone. An important question, therefore, is whether Medicare significantly affects the labor supply of the elderly. This question is crucial when considering Medicare reforms; the fiscal effects of such reforms depend on how labor supply responds. However, there is relatively little research on the labor supply responses to Medicare.

This paper provides an empirical analysis of the effect of employer-provided health insurance and Medicare in determining retirement behavior. Using

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[2]Figures taken from *2010 Medicare Annual Report* (Boards of Trustees of the Federal Hospital Insurance and Federal Supplementary Medical Insurance Trust Funds (2010)).

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data from the Health and Retirement Study (HRS), we estimate a dynamic programming model of retirement that accounts for both saving and uncertain medical expenses. Our results suggest that Medicare is important for understanding retirement behavior, because it insures against medical expense shocks that can exhaust a household's savings.

Our work builds upon, and in part reconciles, several earlier studies. Assuming that individuals value health insurance at the cost paid by employers, Lumsdaine, Stock, and Wise (1994) and Gustman and Steinmeier (1994) found that health insurance has a small effect on retirement behavior. One possible reason for their results is that they found that the average employer contribution to health insurance is modest, and declines by only a small amount after age 65. If workers are risk-averse, however, and if health insurance allows them to smooth consumption when facing volatile medical expenses, they could value employer-provided health insurance well beyond the cost paid by employers. Medicare's age-65 work disincentive thus comes not only from the reduction in average medical costs paid by those without employer-provided health insurance, but also from the reduction in the volatility of those costs.

Addressing this point, Rust and Phelan (1997) and Blau and Gilleskie (2006, 2008) estimated dynamic programming models that account explicitly for risk aversion and uncertainty about out-of-pocket medical expenses. Their estimated labor supply responses to health insurance are larger than those found in studies that omit medical expense risk. Rust and Phelan and Blau and Gilleskie, however, assumed that an individual's consumption equals his income net of out-of-pocket medical expenses. In other words, they ignored an individual's ability to smooth consumption through saving. If individuals can self-insure against medical expense shocks by saving, prohibiting saving will overstate the consumption volatility caused by medical cost volatility. It is therefore likely that Rust and Phelan and Blau and Gilleskie overstated the value of health insurance, and thus the effect of health insurance on retirement.

In this paper we construct a life-cycle model of labor supply that not only accounts for medical expense uncertainty and health insurance, but also has a saving decision. Moreover, we include the coverage provided by means-tested social insurance to account for the fact that Medicaid provides a substitute for other forms of health insurance. To our knowledge, ours is the first study of its kind. While van der Klaauw and Wolpin (2008) and Casanova (2010) also estimated retirement models that account for both savings and uncertain medical expenses, they did not focus on the role of health insurance, and thus use much simpler models of medical expenses.

Almost everyone becomes eligible for Medicare at age 65. However, the Social Security system and pensions also provide retirement incentives at age 65. This makes it difficult to determine whether the high job exit rates observed at age 65 are due to Medicare, Social Security, or pensions. One way we address this problem is to exploit variation in employer-provided health insurance. Some individuals receive employer-provided health insurance only while they work, so that their coverage is tied to their job. Other individuals have retiree coverage, and receive employer-provided health insurance even if they retire. If workers value access to health insurance, those with retiree coverage should be more willing to retire before age 65. Our data show that individuals with retiree coverage tend to retire about $\frac{1}{2}$ year earlier than individuals with tied coverage. This suggests that employer-provided health insurance is a determinant of retirement.

One problem with using employer-provided health insurance to identify Medicare's effect on retirement is that individuals may choose to work for a firm because of its postretirement benefits. The fact that early retirement is common for individuals with retiree coverage may not reflect the effect of health insurance on retirement. Instead, individuals with preferences for early retirement may be self-selecting into jobs that provide retiree coverage. To address this issue, we measure self-selection into jobs with different health insurance plans. We allow the value of leisure and the time discount factor to vary across individuals. Modelling preference heterogeneity with the approach used by Keane and Wolpin (2007), we find that individuals with strong preferences for leisure are more likely to work for firms that provide retiree health insurance. However, self-selection does not affect our main results.

Estimating the model by the method of simulated moments, we find that the model fits the data well with reasonable parameter values. Next, we simulate the labor supply response to changing some of the Medicare and Social Security retirement program rules. Raising the Medicare eligibility age from 65 to 67 would increase years worked by 0.074 years. Eliminating 2 years worth of Social Security benefits would increase years worked by 0.076 years. Thus, even after allowing for both saving and self-selection into health insurance plans, the effect of Medicare on labor supply is as large as the effect of Social Security. One reason why we find that Medicare is important is that we find that medical expense risk is important. Even when we allow individuals to save, they value the consumption smoothing benefits of health insurance. We find that about half the value a typical worker places on his employer-provided health insurance comes from these benefits.

The rest of paper proceeds as follows. Section 2 develops our dynamic programming model of retirement behavior. Section 3 describes how we estimate the model using the method of simulated moments. Section 4 describes the HRS data that we use in our analysis. Section 5 presents life-cycle profiles drawn from these data. Section 6 contains preference parameter estimates for the structural model, and an assessment of the model's performance, both within and outside of the estimation sample. In Section 7, we conduct several policy experiments. In Section 8 we consider a few robustness checks. Section 9 concludes. A set of supplemental appendices comprises the Supplemental Material (French and Jones (2011)) and provides details of our methodology and data, along with additional results.

2. THE MODEL

To capture the richness of retirement incentives, our model is very complex and has many parameters. Appendix A provides definitions for all the variables.

2.1. *Preferences and Demographics*

Consider a household head seeking to maximize his expected discounted (where the subjective discount factor is β) lifetime utility at age t, $t = 59, 60, \ldots, 94$. Each period that he lives, the individual derives utility from consumption, C_t, and hours of leisure, L_t. The within-period utility function is of the form

$$U(C_t, L_t) = \frac{1}{1-\nu}(C_t^{\gamma} L_t^{1-\gamma})^{1-\nu}. \tag{1}$$

We allow both β and γ to vary across individuals. Individuals with higher values of β are more patient, while individuals with higher values of γ place less weight on leisure.

The quantity of leisure is

$$L_t = L - N_t - \phi_{Pt} P_t - \phi_{\mathrm{RE}} \mathrm{RE}_t - \phi_H H_t, \tag{2}$$

where L is the individual's total annual time endowment. Participation in the labor force is denoted by P_t, a 0–1 indicator equal to 1 when hours worked, N_t, are positive. The fixed cost of work, ϕ_{Pt}, is treated as a loss of leisure. Including fixed costs helps us capture the empirical regularity that annual hours of work are clustered around 2,000 hours and 0 hours (Cogan (1981)). Following a number of studies,[3] we allow preferences for leisure, in our case the value of ϕ_{Pt}, to increase linearly with age. Workers who leave the labor force can reenter; reentry is denoted by the 0–1 indicator $\mathrm{RE}_t = 1\{P_t = 1 \text{ and } P_{t-1} = 0\}$, and individuals who reenter the labor market incur cost ϕ_{RE}. The quantity of leisure also depends on an individual's health status through the 0–1 indicator $H_t = 1\{\text{health}_t = \text{bad}\}$, which equals 1 when his health is bad.

Workers alive at age t survive to age $t+1$ with probability s_{t+1}. Following De Nardi (2004), workers who die value bequests of assets, A_t, according to the function

$$b(A_t) = \theta_B \frac{(A_t + \kappa)^{(1-\nu)\gamma}}{1-\nu}. \tag{3}$$

The survival probability s_t, along with the transition probabilities for the health variable H_t, depend on age and previous health status.

[3]Examples include Rust and Phelan (1997), Blau and Gilleskie (2006, 2008), Gustman and Steinmeier (2005), Rust, Buchinsky, and Benitez-Silva (2003), and van der Klaauw and Wolpin (2008).

2.2. *Budget Constraints*

The individual holds three forms of wealth: assets (including housing); pensions; and Social Security. He has several sources of income: asset income, rA_t, where r denotes the constant pre-tax interest rate; labor income, W_tN_t, where W_t denotes wages; spousal income, ys_t; pension benefits, pb_t; Social Security benefits, ss_t; and government transfers, tr_t. The asset accumulation equation is

$$A_{t+1} = A_t + Y_t + \mathrm{ss}_t + \mathrm{tr}_t - M_t - C_t, \tag{4}$$

where M_t denotes medical expenses. Post-tax income, $Y_t = Y(rA_t + W_tN_t + ys_t + \mathrm{pb}_t, \tau)$, is a function of taxable income and the vector τ, described in the Supplemental Material Appendix B, that captures the tax structure.

Individuals face the borrowing constraint

$$A_t + Y_t + \mathrm{ss}_t + \mathrm{tr}_t - C_t \geq 0. \tag{5}$$

Because it is illegal to borrow against future Social Security benefits and difficult to borrow against many forms of future pension benefits, individuals with low nonpension, non-Social Security wealth may not be able to finance their retirement before their Social Security benefits become available at age 62 (Kahn (1988), Rust and Phelan (1997), Gustman and Steinmeier (2005)).[4]

Following Hubbard, Skinner, and Zeldes (1994, 1995), government transfers provide a consumption floor:

$$\mathrm{tr}_t = \max\{0, C_{\min} - (A_t + Y_t + \mathrm{ss}_t)\}. \tag{6}$$

Equation (6) implies that government transfers bridge the gap between an individual's "liquid resources" (the quantity in the inner parentheses) and the consumption floor. Treating $C_{\min}$ as a sustenance level, we further require that $C_t \geq C_{\min}$. Our treatment of government transfers implies that individuals will always consume at least $C_{\min}$, even if their out-of-pocket medical expenses exceed their financial resources.

2.3. *Medical Expenses, Health Insurance, and Medicare*

We define M_t as the sum of all out-of-pocket medical expenses, including insurance premia and expenses covered by the consumption floor. We assume that an individual's medical expenses depend on five components. First,

[4]We assume time-t medical expenses are realized after time-t labor decisions have been made. We view this as preferable to the alternative assumption that the time-t medical expense shocks are fully known when workers decide whether to hold on to their employer-provided health insurance. Given the borrowing constraint and timing of medical expenses, an individual who has extremely high medical expenses this year could have negative net worth next year. Because many people in our data have unresolved medical expenses, medical expense debt seems reasonable.

medical expenses depend on the individual's employer-provided health insurance, I_t. Second, they depend on whether the person is working, P_t, because workers who leave their job often pay a larger fraction of their insurance premiums. Third, they depend on the individual's self-reported health status, H_t. Fourth, medical expenses depend on age. At age 65, individuals become eligible for Medicare, which is a close substitute for employer-provided coverage.[5] Offsetting this, as people age their health declines (in a way not captured by H_t), raising medical expenses. Finally, medical expenses depend on the person-specific component ψ_t, yielding

$$\ln M_t = m(H_t, I_t, t, P_t) + \sigma(H_t, I_t, t, P_t) \times \psi_t. \tag{7}$$

Note that health insurance affects both the expectation of medical expenses, through $m(\cdot)$, and the variance, through $\sigma(\cdot)$

Even after controlling for health status, French and Jones (2004a) found that medical expenses are very volatile and persistent. Thus we model the person-specific component of medical expenses, ψ_t, as

$$\psi_t = \zeta_t + \xi_t, \quad \xi_t \sim N(0, \sigma_\xi^2), \tag{8}$$

$$\zeta_t = \rho_m \zeta_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \sigma_\varepsilon^2), \tag{9}$$

where ξ_t and ε_t are serially and mutually independent; ξ_t is the transitory component, while ζ_t is the persistent component, with autocorrelation ρ_m.

We assume that medical expenditures are exogenous. It is not clear ex ante whether this causes us to understate or overstate the importance of health insurance. On the one hand, individuals who have health insurance receive better care. Our model does not capture this benefit, and in this respect understates the value of health insurance. Conversely, treating medical expenses as exogenous ignores the ability of workers to offset medical shocks by adjusting their expenditures on medical care. This leads us to overstate the consumption risk facing uninsured workers, and thus the value of health insurance. Evidence from other structural analyses suggests that our assumption of exogeneity leads us to overstate the effect of health insurance on retirement.[6]

[5]Individuals who have paid into the Medicare system for at least 10 years become eligible at age 65. A more detailed description of the Medicare eligibility rules is available at http://www.medicare.gov/.

[6]To our knowledge, Blau and Gilleskie (2006) is the only estimated, structural retirement study to have endogenous medical expenditures. Although Blau and Gilleskie (2006) did not discuss how their results would change if medical expenses were treated as exogenous, they found that even with several mechanisms (such as prescription drug benefits) omitted, health insurance has "a modest impact on employment behavior among older males." De Nardi, French, and Jones (2010) studied the saving behavior of retirees. They found that the effects of reducing means-tested social insurance are smaller when medical care is endogenous, rather than exogenous. They also found, however, that even when medical expenditures are a choice variable, they are a major reason why the elderly save.

Differences in labor supply behavior across health insurance categories are an integral part of identifying our model. We assume that there are three mutually exclusive categories of health insurance coverage. The first is *retiree* coverage, where workers keep their health insurance even after leaving their jobs. The second category is *tied* health insurance, where workers receive employer-provided coverage as long as they continue to work. If a worker with tied health insurance leaves his job, he can keep his health insurance coverage for that year. This is meant to proxy for the fact that most firms must provide COBRA health insurance to workers after they leave their job. After 1 year of tied coverage and not working, the individual's insurance ceases.[7] The third category consists of individuals whose potential employers provide no health insurance at all, or *none*. Workers move between these insurance categories according to

$$I_t = \begin{cases} \text{retiree}, & \text{if } I_{t-1} = \text{retiree}, \\ \text{tied}, & \text{if } I_{t-1} = \text{tied and } N_{t-1} > 0, \\ \text{none}, & \text{if } I_{t-1} = \text{none or } (I_{t-1} = \text{tied and } N_{t-1} = 0). \end{cases} \tag{10}$$

2.4. *Wages and Spousal Income*

We assume that the logarithm of wages at time t, $\ln W_t$, is a function of health status (H_t), age (t), hours worked (N_t), and an autoregressive component, ω_t:

$$\ln W_t = W(H_t, t) + \alpha \ln N_t + \omega_t. \tag{11}$$

The inclusion of hours, N_t, in the wage determination equation captures the empirical regularity that, all else equal, part-time workers earn relatively lower wages than full time workers. French (2005) and Erosa, Fuster, and Kambourov (2010) used similar frameworks. The autoregressive component ω_t has the correlation coefficient ρ_W and the normally distributed innovation η_t:

$$\omega_t = \rho_W \omega_{t-1} + \eta_t, \quad \eta_t \sim N(0, \sigma_\eta^2). \tag{12}$$

Because spousal income can serve as insurance against medical shocks, we include it in the model. In the interest of computational simplicity, we assume that spousal income is a deterministic function of an individual's age and health status:

$$ys_t = ys(H_t, t). \tag{13}$$

[7]Although there is some variability across states as to how long individuals are eligible for employer-provided health insurance coverage, by Federal law most individuals are covered for 18 months (Gruber and Madrian (1995)). Given a model period of 1 year, we approximate the 18-month period as 1 year. We do not model the option to take up COBRA, assuming that the take-up rate is 100%. Although the actual take-up rate is around $\frac{2}{3}$ (Gruber and Madrian (1996)), we simulated the model by assuming that the rate was 0%, so that individuals transitioned from tied to none as soon as they stopped working, and found very similar labor supply patterns. Thus assuming a 100% take-up rate does not seem to drive our results.

2.5. *Social Security and Pensions*

Because pensions and Social Security generate potentially important retirement incentives, we model the two programs in detail.

Individuals receive no Social Security benefits until they apply. Individuals can first apply for benefits at age 62. Upon applying, the individual receives benefits until death. The individual's Social Security benefits depend on his average indexed monthly earnings (AIME), which is roughly his average income during his 35 highest earnings years in the labor market.

The Social Security system provides three major retirement incentives.[8] First, while income earned by workers with less than 35 years of earnings automatically increases their AIME, income earned by workers with more than 35 years of earnings increases their AIME only if it exceeds earnings in some previous year of work. Because Social Security benefits increase in AIME, this causes work incentives to drop after 35 years in the labor market. We describe the computation of AIME in more detail in the Supplemental Material Appendix C.

Second, the age at which the individual applies for Social Security affects the level of benefits. For every year before age 65 the individual applies for benefits, benefits are reduced by 6.67% of the age-65 level. This is roughly actuarially fair. But for every year after age 65 that benefit application is delayed, benefits rise by 5.5% up until age 70. This is less than actuarially fair, and encourages people to apply for benefits by age 65.

Third, the Social Security Earnings Test taxes labor income of beneficiaries at a high rate. For individuals aged 62–64, each dollar of labor income above the "test" threshold of $9,120 leads to a 1/2 dollar decrease in Social Security benefits, until all benefits have been taxed away. For individuals aged 65–69 before 2000, each dollar of labor income above a threshold of $14,500 leads to a 1/3 dollar decrease in Social Security benefits, until all benefits have been taxed away. Although benefits taxed away by the earnings test are credited to future benefits, after age 64 the crediting rate is less than actuarially fair, so that the Social Security Earnings Test effectively taxes the labor income of beneficiaries aged 65–69.[9] When combined with the aforementioned incentives to draw Social Security benefits by age 65, the Earnings Test discourages work after age 65. In 2000, the Social Security Earnings Test was abolished for those 65 and older. Because those born in 1933 (the average birth year in our sample)

[8]A description of the Social Security rules can be found in recent editions of the *Green Book* (Committee on Ways and Means). Some of the rules, such as the benefit adjustment formula, depend on an individual's year of birth. Because we fit our model to a group of individuals who on average were born in 1933, we use the benefit formula for that birth year.

[9]The credit rates are based on the benefit adjustment formula. If a year's worth of benefits are taxed away between ages 62 and 64, benefits in the future are increased by 6.67%. If a year's worth of benefits are taxed away between ages 65 and 66, benefits in the future are increased by 5.5%.

turned 67 in 2000, we assume that the earnings test was repealed at age 67. These incentives are incorporated in the calculation of ss_t, which is defined to be net of the earnings test.

Pension benefits, pb_t, are a function of the worker's age and pension wealth. Pension wealth (the present value of pension benefits) in turn depends on pension accruals. We assume that pension accruals are a function of a worker's age, labor income, and health insurance type, using a formula estimated from confidential HRS pension data. The data show that pension accrual rates differ greatly across health insurance categories; accounting for these differences is essential in isolating the effects of employer-provided health insurance. When finding an individual's decision rules, we assume further that the individual's existing pension wealth is a function of his Social Security wealth, age, and health insurance type. Details of our pension model are described in Section 4.3 and Supplemental Material Appendix D.

2.6. *Recursive Formulation*

In addition to choosing hours and consumption, eligible individuals decide whether to apply for Social Security benefits; let the indicator variable $B_t \in \{0, 1\}$ equal 1 if an individual has applied. In recursive form, the individual's problem can be written as

$$
\begin{aligned}
(14)\qquad V_t(X_t) = \max_{C_t, N_t, B_t} \bigg\{ & \frac{1}{1-\nu}(C_t^{\gamma}(L - N_t - \phi_{Pt}P_t \\
& - \phi_{\mathrm{RE}}\mathrm{RE}_t - \phi_H H_t)^{1-\gamma})^{1-\nu} \\
& + \beta(1 - s_{t+1})b(A_{t+1}) \\
& + \beta s_{t+1} \int V_{t+1}(X_{t+1})\, dF(X_{t+1}|X_t, t, C_t, N_t, B_t) \bigg\},
\end{aligned}
$$

subject to equations (5) and (6). The vector $X_t = (A_t, B_{t-1}, H_t, \mathrm{AIME}_t, I_t, P_{t-1}, \omega_t, \zeta_{t-1})$ contains the individual's state variables, while the function $F(\cdot|\cdot)$ gives the conditional distribution of these state variables, using equations (4) and (7)–(13).[10] The solution to the individual's problem consists of the consumption rules, work rules, and benefit application rules that solve equation (14). These decision rules are found numerically using value function iteration. Supplemental Material Appendix E describes our numerical methodology.

[10]Spousal income and pension benefits (see Supplemental Material Appendix D) depend only on the other state variables and are thus not state variables themselves.

3. ESTIMATION

To estimate the model, we adopt a two-step strategy, similar to the one used by Gourinchas and Parker (2002), French (2005), and Laibson, Repetto, and Tobacman (2007). In the first step, we estimate or calibrate parameters that can be cleanly identified without explicitly using our model. For example, we estimate mortality rates and health transitions straight from demographic data. In the second step, we estimate the preference parameters of the model, along with the consumption floor, using the method of simulated moments (MSM).[11]

3.1. *Moment Conditions*

The objective of MSM estimation is to find the preference vector that yields simulated life-cycle decision profiles that "best match" (as measured by a GMM criterion function) the profiles from the data. The following moment conditions comprise our estimator:

(i) Because an individual's ability to self-insure against medical expense shocks depends on his asset level, we match $1/3$ and $2/3$ asset quantiles by age. We match these quantiles in each of T periods (ages), for a total of $2T$ moment conditions.

(ii) We match job exit rates by age for each health insurance category. With three health insurance categories (none, retiree, and tied), this generates $3T$ moment conditions.

(iii) Because the value a worker places on employer-provided health insurance may depend on his wealth, we match labor force participation conditional on the combination of asset quantile and health insurance status. With two quantiles (generating three quantile-conditional means) and three health insurance types, this generates $9T$ moment conditions.

(iv) To help identify preference heterogeneity, we utilize a series of questions in the HRS that ask workers about their preferences for work. We combine the answers to these questions into a time-invariant index, $\text{pref} \in \{\text{high}, \text{low}, \text{out}\}$, which is described in greater detail in Section 4.4. Matching participation conditional on each value of this index generates another $3T$ moment conditions.

(v) Finally, we match hours of work and participation conditional on our binary health indicator. This generates $4T$ moment conditions.

Combined, the five preceding items result in $21T$ moment conditions. Supplemental Material Appendix F provides a detailed description of the moment conditions, the mechanics of our MSM estimator, the asymptotic distribution of our parameter estimates, and our choice of weighting matrix.

[11] An early application of the MSM to a structural retirement model is Berkovec and Stern (1991).

3.2. *Initial Conditions and Preference Heterogeneity*

A key part of our estimation strategy is to compare the behavior of individuals with different forms of employer-provided health insurance. If access to health insurance is an important factor in the retirement decision, we should find that individuals with tied coverage retire later than those with retiree coverage. In making such a comparison, however, we must account for the possibility that individuals with different health insurance options differ systematically along other dimensions as well. For example, individuals with retiree coverage tend to have higher wages and more generous pensions.

We control for this "initial conditions" problem in three ways. First, the initial distribution of simulated individuals is drawn directly from the data. Because households with retiree coverage are more likely to be wealthy in the data, households with retiree coverage are more likely to be wealthy in our initial distribution. Similarly, in our initial distribution, households with the high levels of education are more likely to have high values of the persistent wage shock ω_t.

Second, we model carefully the way in which pension and Social Security accrual varies across individuals and groups.

Finally, we control for unobservable differences across health insurance groups by introducing permanent preference heterogeneity, using the approach introduced by Heckman and Singer (1984) and adapted by (among others) Keane and Wolpin (1997) and van der Klaauw and Wolpin (2008). Each individual is assumed to belong to one of a finite number of preference "types," with the probability of belonging to a particular type a logistic function of the individual's initial state vector: his age, wealth, initial wages, health status, health insurance type, medical expenditures, and preference index.[12] We estimate the type probability parameters jointly with the preference parameters and the consumption floor.

In our framework, correlations between preferences and health insurance emerge because people with different preferences systematically select jobs with different types of health insurance coverage. Workers in our data set are first observed in their 50s; by this age, all else equal, jobs that provide generous postretirement health insurance are more likely to be held by workers who wish to retire early. One way to measure this self-selection is to structurally model the choice of health insurance at younger ages, and use the predictions of that

[12]These discrete type-based differences are the only preference heterogeneity in our model. For this reason many individuals in the data make decisions different from what the model would predict. Our MSM procedure circumvents this problem by using moment conditions that average across many individuals. One way to reconcile model predictions with individual observations is to introduce measurement error. In earlier drafts of this paper (French and Jones (2004b)) we considered this possibility by estimating a specification where we allowed for measurement error in assets. Adding measurement error, however, had little effect on either the preference parameter estimates or policy experiments, and we dropped this case.

model to infer the correlation between preferences and health insurance in the first wave of the HRS. Because such an approach is computationally expensive, we instead model the correlation between preferences and health insurance in the initial conditions.

3.3. *Wage Selection*

We estimate a selection-adjusted wage profile using the procedure developed in French (2005). First, we estimate a fixed-effects wage profile from HRS data, using the wages observed for individuals who are working. The fixed-effects estimator is identified using wage growth for workers. If wage growth rates for workers and nonworkers are the same, composition bias problems—the question of whether high wage individuals drop out of the labor market later than low wage individuals—are not a problem. However, if individuals leave the market because of a wage drop, such as from job loss, then wage growth rates for workers will be greater than wage growth for nonworkers. This selection problem will bias estimated wage growth upward.

We control for selection bias by finding the wage profile that, when fed into our model, generates the same fixed-effects profile as the HRS data. Because the simulated fixed-effect profiles are computed using only the wages of those simulated agents that work, the profiles should be biased upward for the same reasons they are in the data. We find this bias-adjusted wage profile using the iterative procedure described in French (2005).

4. DATA AND CALIBRATIONS

4.1. *HRS Data*

We estimate the model using data from the Health and Retirement Survey (HRS). The HRS is a sample of noninstitutionalized individuals, aged 51–61 in 1992, and their spouses. With the exception of assets and medical expenses, which are measured at the household level, our data are for male household heads. The HRS surveys individuals every 2 years, so that we have eight waves of data covering the period 1992–2006. The HRS also asks respondents retrospective questions about their work history that allow us to infer whether the individual worked in nonsurvey years. Details of this, as well as variable definitions, selection criteria, and a description of the initial joint distribution, are in Supplemental Material Appendix G.

As noted above, the Social Security rules depend on an individual's year of birth. To ensure that workers in our sample face a similar set of Social Security retirement rules, we fit our model to the data for the cohort of individuals aged 57–61 in 1992. However, when estimating the stochastic processes that individuals face, we use the full sample, plus Assets and Health Dynamics of the Oldest Old (AHEAD) data, which provides information on these processes at older ages. With the exception of wages, we do not adjust the data for cohort effects. Because our subsample of the HRS covers a fairly narrow age range, this omission should not generate much bias.

4.2. *Health Insurance and Medical Expenses*

We assign individuals to one of three mutually exclusive health insurance groups: retiree, tied, and none, as described in Section 2. Because of small sample problems, the none group includes those who have private health insurance as well as those who have no insurance at all. Both face high medical expenses because they lack employer-provided coverage. Private health insurance is a poor substitute for employer-provided coverage, as high administrative costs and adverse selection problems can result in prohibitively expensive premiums. Moreover, private insurance is much less likely to cover preexisting medical conditions. Because the model includes a consumption floor to capture the insurance provided by Medicaid, the none group also includes those who receive health care through Medicaid. We assign those who have health insurance provided by their spouse to the retiree group, along with those who report that they could keep their health insurance if they left their jobs. Both of these groups have health insurance that is not tied to their job. We assign individuals who would lose their employer-provided health insurance after leaving their job to the tied group. Supplemental Material Appendix H shows our estimated (health insurance-conditional) job exit rate profiles are robust to alternative coding decisions.

The HRS has data on self-reported medical expenses. Medical expenses are the sum of insurance premia paid by households, drug costs, and out-of-pocket costs for hospital, nursing home care, doctor visits, dental visits, and outpatient care. Because our model explicitly accounts for government transfers, the appropriate measure of medical expenses includes expenses paid for by government transfers. Unfortunately, we observe only the medical expenses paid by households, not those paid by Medicaid. Therefore, we impute Medicaid payments for households that received Medicaid benefits, as described in Supplemental Material Appendix G.

We fit these data to the medical expense model described in Section 2. Because of small sample problems, we allow the mean, $m(\cdot)$, and standard deviation, $\sigma(\cdot)$, to depend only on the individual's Medicare eligibility, health insurance type, health status, labor force participation, and age. Following the procedure described in French and Jones (2004a), $m(\cdot)$ and $\sigma(\cdot)$ are set so that the model replicates the mean and 95th percentile of the cross-sectional distribution of medical expenses in each of these categories. Details are provided in Supplemental Material Appendix I.

Table I presents summary statistics (in 1998 dollars), conditional on health status. Table I shows that for healthy individuals who are 64 years old, and thus not receiving Medicare, average annual medical expenses are \$3,360 for workers with tied coverage and \$6,010 for those with none, a difference of

TABLE I

MEDICAL EXPENSES, BY MEDICARE AND HEALTH INSURANCE STATUS

	Retiree		Tied		
	Working	Not Working	Working	Not Working	None
Age = 64, without Medicare, Good Health					
Mean	\$3,160	\$3,880	\$3,360	\$5,410	\$6,010
Standard deviation	\$5,460	\$7,510	\$5,040	\$10,820	\$15,830
99.5th percentile	\$32,700	\$44,300	\$30,600	\$63,500	\$86,900
Age = 65, with Medicare, Good Health					
Mean	\$3,320	\$3,680	\$3,830	\$4,230	\$4,860
Standard deviation	\$4,740	\$5,590	\$5,920	\$9,140	\$7,080
99.5th percentile	\$28,800	\$33,900	\$35,800	\$52,800	\$43,000
Age = 64, without Medicare, Bad Health					
Mean	\$3,930	\$4,830	\$4,170	\$6,730	\$7,470
Standard deviation	\$6,940	\$9,530	\$6,420	\$13,740	\$20,060
99.5th percentile	\$41,500	\$56,100	\$38,900	\$80,400	\$109,500
Age = 65, with Medicare, Bad Health					
Mean	\$4,130	\$4,580	\$4,760	\$5,260	\$6,040
Standard deviation	\$6,030	\$7,120	\$7,530	\$11,590	\$9,020
99.5th percentile	\$36,600	\$43,000	\$45,500	\$66,700	\$54,700

\$2,650. With the onset of Medicare at age 65, the difference shrinks to \$1,030.[13] Thus, the value of having employer-provided health insurance coverage largely vanishes at age 65.

As Rust and Phelan (1997) emphasized, it is not just differences in mean medical expenses that determine the value of health insurance, but also differences in variance and skewness. If health insurance reduces medical expense volatility, risk-averse individuals may value health insurance at well beyond the cost paid by employers. To give a sense of the volatility, Table I also presents the standard deviation and 99.5th percentile of the medical expense distributions. Table I shows that for healthy individuals who are 64 years old, annual medical expenses have a standard deviation of \$5,040 for workers with tied coverage and \$15,830 for those with none, a difference of \$10,790. With the onset of Medicare at age 65, the difference shrinks to \$1,160. Therefore, Medicare not only reduces average medical expenses for those without employer-provided health insurance, it reduces medical expense volatility as well.

The parameters for the idiosyncratic process ψ_t, $(\sigma_\xi^2, \sigma_\varepsilon^2, \rho_m)$, are taken from French and Jones (2004a, "fitted" specification). Table II presents the parame-

[13]The pre-Medicare cost differences are roughly comparable to The Employee Benefit Research Institute (EBRI) (1999) estimated that employers on average contribute \$3,288 per year to their employees' health insurance. They are larger than Gustman and Steinmeier's (1994) estimate that employers contribute about \$2,500 per year before age 65 (1977 NMES data, adjusted to 1998 dollars with the medical component of the consumer price index (CPI)).

TABLE II

VARIANCE AND PERSISTENCE OF INNOVATIONS TO MEDICAL EXPENSES

Parameter	Variable	Estimate (Standard Errors)
ρ_m	Autocorrelation of persistent component	0.925 (0.003)
σ_ε^2	Innovation variance of persistent component	0.04811 (0.008)
σ_ξ^2	Innovation variance of transitory component	0.6668 (0.014)

ters, which have been normalized so that the overall variance, σ_ψ^2, is 1. Table II reveals that at any point in time, the transitory component generates almost 67% of the cross-sectional variance in medical expenses. The results in French and Jones reveal, however, that most of the variance in cumulative *lifetime* medical expenses is generated by innovations to the persistent component. For this reason, the cross-sectional distribution of medical expenses reported in Table I understates the lifetime risk of medical expenses. Given the autocorrelation coefficient ρ_m of 0.925, this is not surprising.

4.3. *Pension Accrual*

Supplemental Material Appendix D describes how we use confidential HRS pension data to construct the accrual rate formula. Figure 1 shows the average pension accrual rates generated by this formula when we simulate the model.

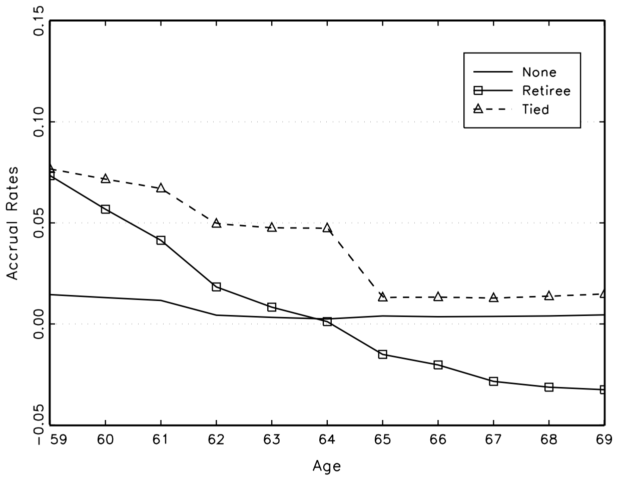

FIGURE 1.—Average pension accrual rates by age and health insurance coverage.

Figure 1 reveals that workers with retiree coverage face the sharpest drops in pension accrual after age 60.[14] While retiree coverage in and of itself provides an incentive for early retirement, the pension plans associated with retiree coverage also provide the strongest incentives for early retirement. Failing to capture this link will lead the econometrician to overstate the effect of retiree coverage on retirement.

4.4. *Preference Index*

To better measure preference heterogeneity in the population (and how it is correlated with health insurance), we estimate a person's "willingness" to work using three questions from the first (1992) wave of the HRS. The first question asks the respondent the extent to which he agrees with the statement, "Even if I didn't need the money, I would probably keep on working." The second question asks the respondent, "When you think about the time when you will retire, are you looking forward to it, are you uneasy about it, or what?" The third question asks, "How much do you enjoy your job?"

To combine these three questions into a single index, we regress waves 5–7 (survey years 2000–2004) participation on the response to the three questions along with polynomials and interactions of all the state variables in the model: age, health status, wages, wealth, AIME, medical expenses, and health insurance type. Multiplying the numerical responses to the three questions by their respective estimated coefficients and summing yields an index. We then discretize the index into three values: high for the top 50% of the index for those working in wave 1; low for the bottom 50% of the index for those working in wave 1; out for those not working in wave 1. Supplemental Material Appendix J provides additional details on the construction of the index. Figure 6 below shows that the index has great predictive power: at age 65, participation rates are 56% for those with an index of high, 39% for those with an index of low, and 12% for those with an index of out.

4.5. *Wages*

Recall from equation (11) that $\ln W_t = \alpha \ln(N_t) + W(H_t, t) + \omega_t$. Following Aaronson and French (2004), we set $\alpha = 0.415$, which implies that a 50% drop in work hours leads to a 25% drop in the offered hourly wage. This is in the middle of the range of estimates of the effect of hours worked on the offered hourly wage.

We estimate $W(H_t, t)$ using the methodology described in Section 3.3.

[14]Because Figure 1 is based on our estimation sample, it does not show accrual rates for earlier ages. Estimates that include the validation sample show, however, that those with retiree coverage have the highest pension accrual rates in their early and middle 50s.

The parameters for the idiosyncratic process ω_t, (σ_η^2, ρ_W), are estimated by French (2005). The results indicate that the autocorrelation coefficient ρ_W is 0.977; wages are almost a random walk. The estimate of the innovation variance σ_η^2 is 0.0141; 1 standard deviation of an innovation in the wage is 12% of wages.

4.6. *Remaining Calibrations*

We set the interest rate r equal to 0.03. Spousal income depends on an age polynomial and health status. Health status and mortality both depend on previous health status interacted with an age polynomial.

5. DATA PROFILES AND INITIAL CONDITIONS

5.1. *Data Profiles*

Figure 2 presents some of the labor market behavior we want our model to explain. The top panel of Figure 2 shows empirical job exit rates by health insurance type. Recall that Medicare should provide the largest labor market incentives for workers who have tied health insurance. If these people place a high value on employer-provided health insurance, they should either work until age 65, when they are eligible for Medicare, or they should work until age 63.5 and use COBRA coverage as a bridge to Medicare. The job exit profiles provide some evidence that those who have tied coverage do tend to work until age 65. While the age-65 job exit rate is similar for those whose health insurance type is tied (20%), retiree (17%), or none (18%), those with retiree coverage have higher exit rates at 62 (22%) than those with tied (14%) or none (18%).[15] At almost every age other than 65, those who have retiree coverage have higher job exit rates than those with tied or no coverage. These differences across health insurance groups, while large, are smaller than the differences in the empirical exit profiles reported by Rust and Phelan (1997).

The low job exit rates before age 65 and the relatively high job exit rates at age 65 for those who have tied coverage suggests that some people who have tied coverage are working until age 65, when they become eligible for Medicare. On the other hand, job exit rates for those who have tied coverage are lower than those who have retiree coverage for every age other than 65, and are not much higher at age 65. This suggests that differences in health insurance coverage may not be the only reason for the differences in job exit rates.

[15] The differences across groups are statistically different at 62, but not at 65. Furthermore, F-tests reject the hypothesis that the three groups have identical exit rates at all ages at the 5% level.

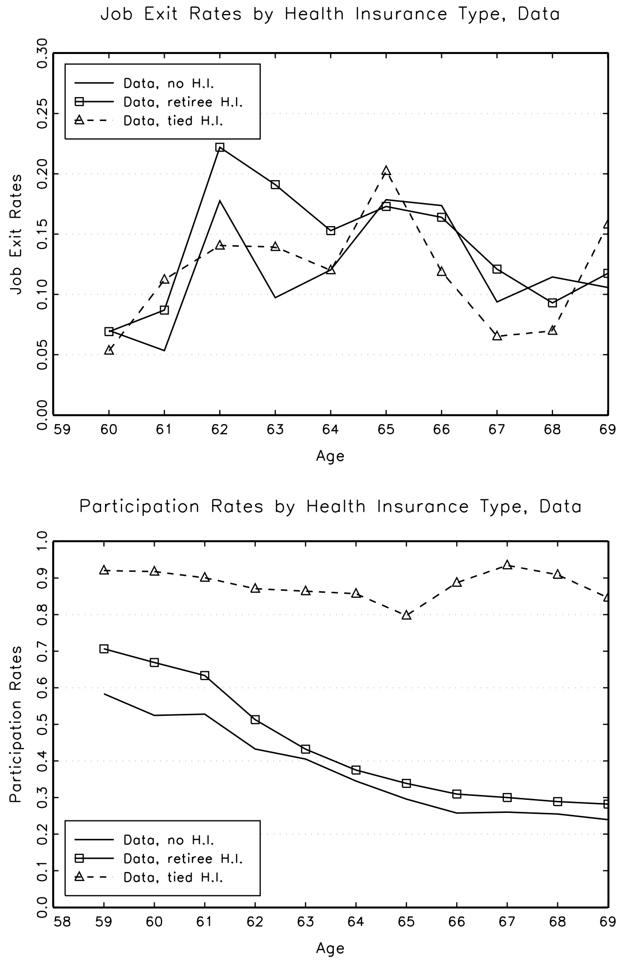

FIGURE 2.—Job exit and participation rates: data.

The bottom panel of Figure 2 presents observed labor force participation rates. In comparing participation rates across health insurance categories, it is useful to keep in mind the transitions implied by equation (10): retiring workers in the tied insurance category transition into the none category. Because of this, the labor force participation rates for those who have tied insurance are calculated for a group of individuals who were all working in the previous period. It is thus not surprising that the tied category has the highest participation rates. Conversely, it is not surprising that the none category has the lowest participation rates, given that it includes tied workers who retire.

5.2. *Initial Conditions*

Each artificial individual in our model begins its simulated life with the year-1992 state vector of an individual, aged 57–61 in 1992, observed in the data. Table III summarizes this initial distribution, the construction of which is described in Supplemental Material Appendix G. Table III shows that individuals with retiree coverage tend to have the most asset and pension wealth, while individuals in the none category have the least. The median individual in the none category has no pension wealth at all. Individuals in the none category are also more likely to be in bad health and, not surprisingly, less likely to be working. In contrast, individuals who have tied coverage have high values of the preference index, suggesting that their delayed retirement reflects differences in preferences as well as in incentives.

TABLE III

SUMMARY STATISTICS FOR THE INITIAL DISTRIBUTION

	Retiree	Tied	None
Age			
Mean	58.7	58.6	58.7
Standard deviation	1.5	1.5	1.5
AIME (in thousands of 1998 dollars)			
Mean	24.9	24.9	16.0
Median	27.2	26.9	16.2
Standard deviation	9.1	8.6	9.0
Assets (in thousands of 1998 dollars)			
Mean	231	205	203
Median	147	118	53
Standard deviation	248	251	307
Pension wealth (in thousands of 1998 dollars)			
Mean	129	80	17
Median	62	17	0
Standard deviation	180	212	102
Wage (in 1998 dollars)			
Mean	17.4	17.6	12.0
Median	14.7	14.6	8.6
Standard deviation	13.4	12.4	11.2
Preference index			
Fraction out	0.27	0.04	0.48
Fraction low	0.42	0.44	0.19
Fraction high	0.32	0.52	0.33
Fraction in bad health	0.20	0.13	0.41
Fraction working	0.73	0.96	0.52
Number of observations	1,022	225	455

6. BASELINE RESULTS

6.1. *Preference Parameter Estimates*

The goal of our MSM estimation procedure is to match the life-cycle profiles for assets, hours, and participation found in the HRS data. To use these profiles to identify preferences, we make several identifying assumptions, the most important being that preferences vary with age in two specific ways: (i) through changes in health status and (ii) through the linear time trend in the fixed cost ϕ_{Pt}. Therefore, age can be thought of as an "exclusion restriction," which changes the incentives for work and savings in ways that cannot be captured with changes in preferences.

Table IV presents preference parameter estimates. The first three rows of Table IV show the parameters that vary across the preference types. We assume that there are three types of individuals, and that the types differ in the utility weight on consumption, γ, and their time discount factor, β. Individuals who have high values of γ have stronger preferences for work; individuals

TABLE IV

ESTIMATED STRUCTURAL PARAMETERS[a]

Parameters That Vary Across Individuals	Type 0	Type 1	Type 2
γ: Consumption weight	0.412 (0.045)	0.649 (0.007)	0.967 (0.203)
β: Time discount factor	0.945 (0.074)	0.859 (0.013)	1.124 (0.328)
Fraction of individuals	0.267	0.615	0.118

Parameters That Are Common to All Individuals			
ν: Coefficient of relative risk aversion, utility	7.49 (0.312)	θ_B: Bequest weight[b]	0.0223 (0.0012)
κ: Bequest shifter, in thousands	444 (28.2)	$c_{\min}$: Consumption floor	4,380 (167)
L: Leisure endowment, in hours	4,060 (44)	ϕ_H: Hours of in hours bad health	506 (20.9)
ϕ_{P0}: Fixed cost of work at age 60, in hours	826 (20.0)	ϕ_{P1}: Fixed cost of work: age trend, in hours	54.7 (2.57)
ϕ_{RE}: Hours of leisure lost when reentering labor market	94.0 (8.64)		
χ^2 Statistic = 751		Degrees of freedom = 171	

[a]Method of simulated moments estimates. Estimates use a diagonal weighting matrix (see Supplemental Material Appendix F for details). Standard errors are given in parentheses. Parameters are estimated jointly with type prediction equation. The estimated coefficients for the type prediction equation are shown in Supplemental Material Appendix K.

[b]Parameter expressed as the marginal propensity to consume out of final-period wealth.

who have high values of β are more patient and thus more willing to defer consumption and leisure. Table IV reveals significant differences in γ and β across preference types, which are discussed in some detail in Section 6.2.

Table IV also shows the fraction of workers who belong to each preference type; the coefficients for the preference type prediction equation are shown in Supplemental Material Appendix K. Averaging over the three types reveals that the average value of β, the discount factor, implied by our model is 0.913, which is slightly lower than most estimates. The discount factor is identified by the intertemporal substitution of consumption and leisure, as embodied in the asset and labor supply profiles.

Another key parameter is ν, the coefficient of relative risk aversion for the consumption–leisure composite. A more familiar measure of risk aversion is the coefficient of relative risk aversion for consumption. Assuming that labor supply is fixed, it can be approximated as $-\frac{(\partial^2 U/\partial C^2)C}{\partial U/\partial C} = -(\gamma(1-\nu)-1)$. The weighted average value of the coefficient is 5.0. This value falls within the range of estimates found in recent studies by Cagetti (2003) and French (2005), but it is larger than the values of 1.1, 1.8, and 1.0 reported by Rust and Phelan (1997), Blau and Gilleskie (2006), and Blau and Gilleskie (2008), respectively, in their studies of retirement.

The risk coefficient ν and the consumption floor $C_{\min}$ are identified in large part by the asset quantiles, which reflect precautionary motives. The bottom quantile in particular depends on the interaction of precautionary motives and the consumption floor. If the consumption floor is sufficiently low, the risk of a catastrophic medical expense shock, which over a lifetime could equal over \$100,000 (see French and Jones (2004a)), will generate strong precautionary incentives. Conversely, as emphasized by Hubbard, Skinner, and Zeldes (1995), a high consumption floor discourages saving among the poor, since the consumption floor effectively imposes a 100% tax on the savings of those with high medical expenses and low income and assets.

Our estimated consumption floor of \$4,380 is similar to other estimates of social insurance transfers for the indigent. For example, when we use Hubbard, Skinner, and Zeldes's (1994, Appendix A) procedures and more recent data, we find that the average benefit available to a childless household with no members aged 65 or older was \$3,500. A value of \$3,500 understates the benefits available to individuals over age 65; in 1998, the Federal SSI benefit for elderly (65+) couples was nearly \$9,000 (Committee on Ways and Means (2000, p. 229)).[16] On the other hand, about half of eligible households do not collect SSI benefits (Elder and Powers (2006, Table 2)), possibly because transactions or "stigma" costs outweigh the value of public assistance. Low take-up

[16]Our framework also lacks explicit disability insurance. A recent structural analysis of this program is Low and Pistaferri (2010).

rates, along with the costs that probably underlie them, suggest that the effective consumption floor need not equal statutory benefits.

The bequest parameters θ_B and κ are identified largely from the top asset quantile. It follows from equation (3) that when the shift parameter κ is large, the marginal utility of bequests will be lower than the marginal utility of consumption unless the individual is rich. In other words, the bequest motive mainly affects the saving of the rich; for more on this point, see De Nardi (2004). Our estimate of θ_B implies that the marginal propensity to consume out of wealth in the final period of life (which is a nonlinear function of θ_B, β, γ, ν, and κ) is 1 for low income individuals and 0.022 for high income individuals.

Turning to labor supply, we find that individuals in our sample are willing to intertemporally substitute their work hours. In particular, simulating the effects of a 2% wage change reveals that the wage elasticity of average hours is 0.486 at age 60. This relatively high labor supply elasticity arises because the fixed cost of work generates volatility on the participation margin. The participation elasticity is 0.353 at age 60, implying that wage changes cause relatively small hours changes for workers. For example, the Frisch labor supply elasticity of a type-1 individual working 2,000 hours per year at age 60 is approximated as $-\frac{L-N_t-\phi_{P0}}{N_t} \times \frac{1}{(1-\gamma)(1-\nu)-1} = 0.19$.

The fixed cost of work at age 60, ϕ_{P0}, is 826 hours per year, and it increases by $\phi_{P1} = 55$ hours per year. The fixed cost of work is identified by the life-cycle profile of hours worked by workers. Average hours of work (available upon request) do not drop below 1,000 hours per year (or 20 hours per week, 50 weeks per year) even though labor force participation rates decline to near zero. In the absence of a fixed cost of work, one would expect hours worked to parallel the decline in labor force participation (Rogerson and Wallenius (2009)). The time endowment L is identified by the combination of participation and hours profiles. The time cost of bad health, ϕ_H, is identified by noting that unhealthy individuals work fewer hours than healthy individuals, even after conditioning on wages. The reentry cost, ϕ_{RE}, of 94 hours is identified by exit rates. In the absence of a reentry cost, workers are more willing to "churn" in and out of the labor force, raising exit rates.

6.2. *Preference Heterogeneity and Health Insurance*

Table IV shows considerable heterogeneity in preferences. To understand these differences, Table V shows simulated summary statistics for each of the preference types. Table V reveals that type-0 individuals have the lowest value of γ, that is, they place the highest value on leisure: 92% of type-0 individuals were out of the labor force in wave 1. Type-2 individuals, in contrast, have the highest value of γ: 84% of type-2 individuals have a preference index of high, meaning that they were working in wave 1 and self-reported having a low preference for leisure. Type-1 individuals fall in the middle, valuing leisure

TABLE V

MEAN VALUES BY PREFERENCE TYPE: SIMULATIONS

	Type 0	Type 1	Type 2
Key Preference Parameters			
γ[a]	0.412	0.649	0.967
β[a]	0.945	0.859	1.124
Means by Preference Type			
Assets ($1,000s)	150	215	405
Pension Wealth ($1,000s)	92	97	74
Wages ($/hour)	11.3	19.0	11.1
Probability of Health Insurance Type, Given Preference Type			
Health insurance = none	0.371	0.222	0.261
Health insurance = retiree	0.607	0.603	0.581
Health insurance = tied	0.023	0.175	0.158
Probability of Preference Index Value, Given Preference Type			
Preference index = out	0.922	0.068	0.034
Preference index = low	0.039	0.539	0.131
Preference index = high	0.039	0.392	0.835
Fraction of individuals	0.267	0.615	0.118

[a]Values of β and γ are from Table IV.

less than type-0 individuals, but more than type-2 individuals: 54% of type-1 individuals have a preference index value of low.

Including preference heterogeneity allows us to control for the possibility that workers with different preferences select jobs with different health insurance packages. Table V suggests that some self-selection is occurring, as it reveals that while 14% of workers with tied coverage are type-2 agents, who have the lowest disutility of work, only 5% are type-0 agents, who have the highest disutility. In contrast, 11% of workers with retiree coverage are type-2 agents and 27% are type-0 agents. This suggests that workers who have tied coverage might be more willing to retire later than those who have retiree coverage because they have a lower disutility of work. However, Section 6.4 shows that accounting for this correlation has little impact on the estimated effect of health insurance on retirement.

6.3. *Simulated Profiles*

The bottom of Table IV displays the overidentification test statistic. Even though the model is formally rejected, the life-cycle profiles generated by the model match up well with the life-cycle profiles found in the data.

Figure 3 shows the 1/3 and 2/3 asset quantiles at each age for the HRS sample and for the model simulations. For example, at age 64 about 1/3 of the men in our sample live in households with less than $80,000 in assets, and about 1/3 live in households with over $270,000 of assets. Figure 3 shows that the

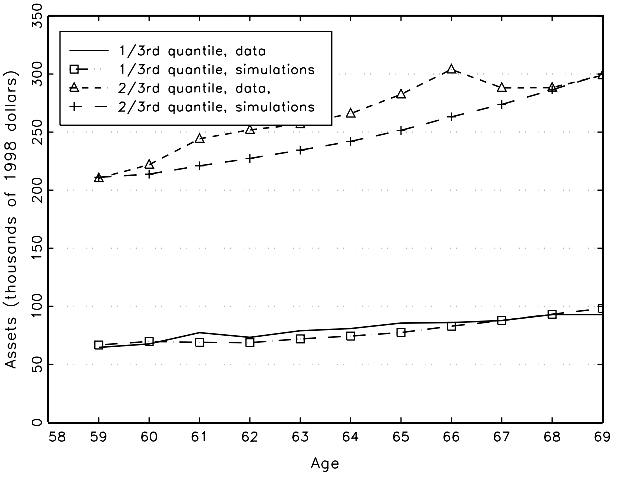

FIGURE 3.—Asset quantiles: data and simulations.

model fits both asset quantiles well. The model is able to fit the lower quantile in large part because of the consumption floor of $4,350; the predicted 1/3 quantile rises when the consumption floor is lowered.

The three panels in the left hand column of Figure 4 show that the model is able to replicate the two key features of how labor force participation varies with age and health insurance. The first key feature is that participation declines with age, and the declines are especially sharp between ages 62 and 65. The model underpredicts the decline in participation at age 65 (a 4.9 percentage point decline in the data versus a 3.5 percentage point decline predicted by the model), but comes closer at age 62 (a 10.6 percentage point decline in the data versus a 10.9 percentage point decline predicted by the model).

The second key feature is that there are large differences in participation and job exit rates across health insurance types. The model does a good job of replicating observed differences in participation rates. For example, the model matches the low participation levels of the uninsured. Turning to the lower left panel of Figure 5, the data show that the group with the lowest participation rates are the uninsured with low assets. The model is able to replicate this fact because of the consumption floor. Without a high consumption floor, the risk of catastrophic medical expenses, in combination with risk aversion, would cause the uninsured to remain in the labor force and accumulate a buffer stock of assets.

The panels in the right hand column of Figure 4 compare observed and simulated job exit rates for each health insurance type. The model does a good job of fitting the exit rates of workers with retiree or tied coverage. For example, the model captures the high age-62 job exit rates for those with retiree coverage and the high age-65 job exit rates for those with tied coverage. How-

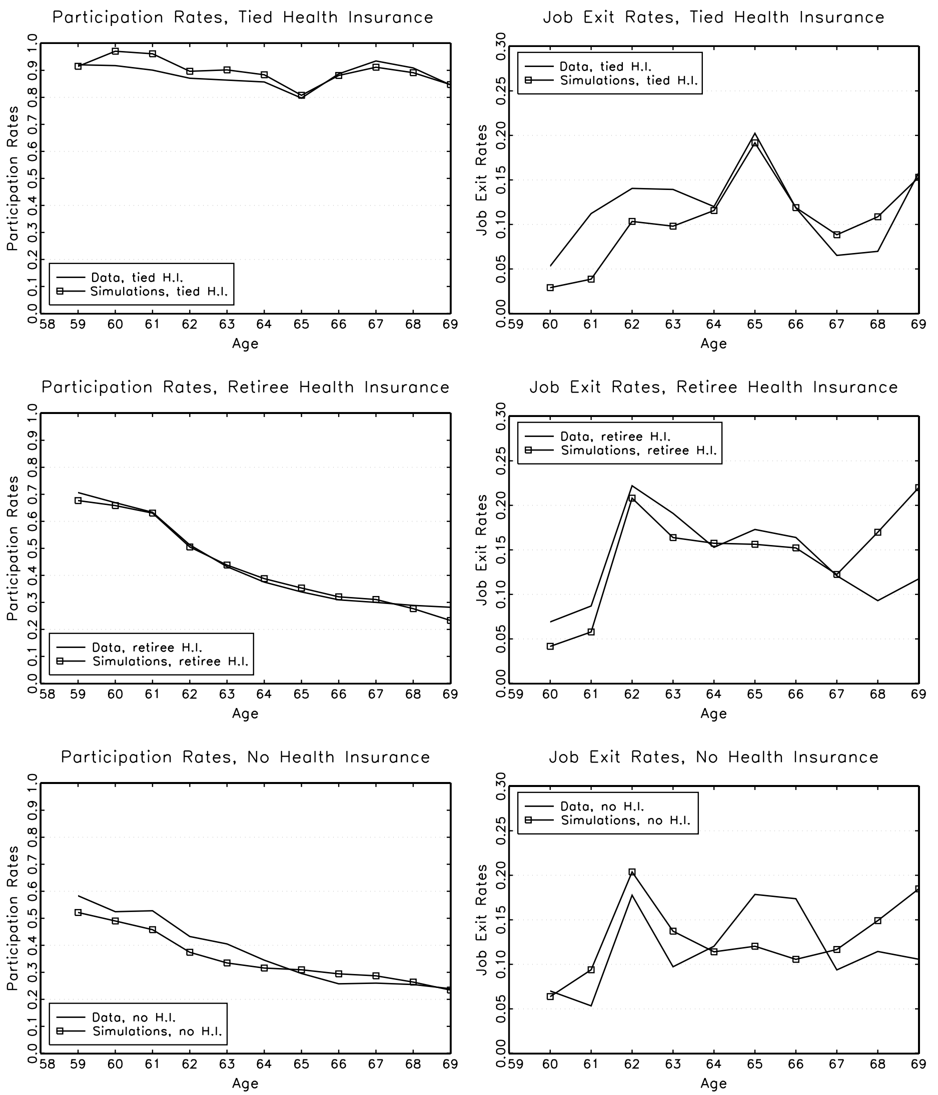

FIGURE 4.—Participation and job exit rates: data and simulations.

ever, it fails to capture the high exit rates at age 65 for workers with no health insurance.

Figure 6 shows how participation differs across the three values of the discretized preference index constructed from HRS attitudinal questions. Recall that an index value of out implies that the individual was not working in 1992. Not surprisingly, participation for this group is always low. Individuals who

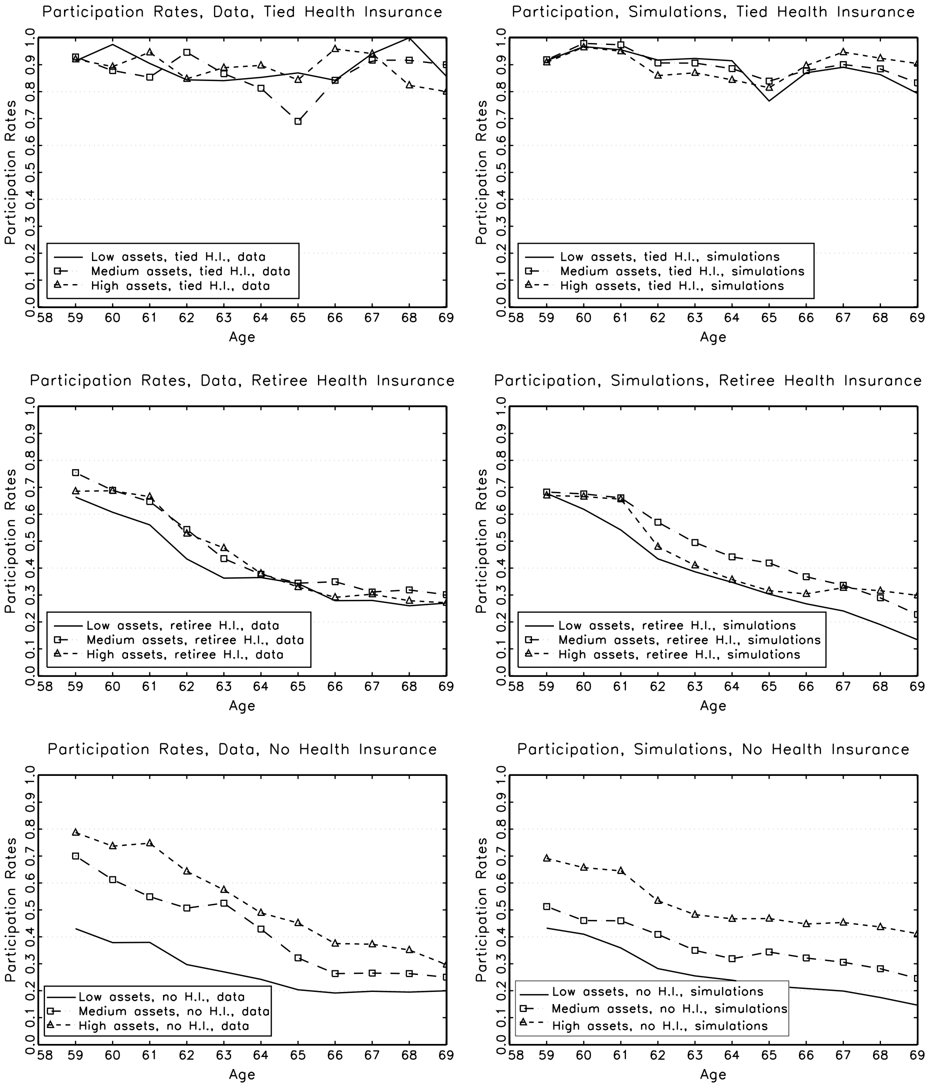

FIGURE 5.—Labor force participation rates by asset grouping: data and simulations.

have positive values of the preference index differ primarily in the rate at which they leave the labor force. Although low-index individuals initially work as much as high-index individuals, they leave the labor force more quickly. As noted in our discussion of the preference parameters, the model replicates these differences by allowing the taste for leisure (γ) and the discount rate (β) to vary across preference types. When we do not allow for preference heterogeneity, the model is unable to replicate the patterns observed in Figure 6.

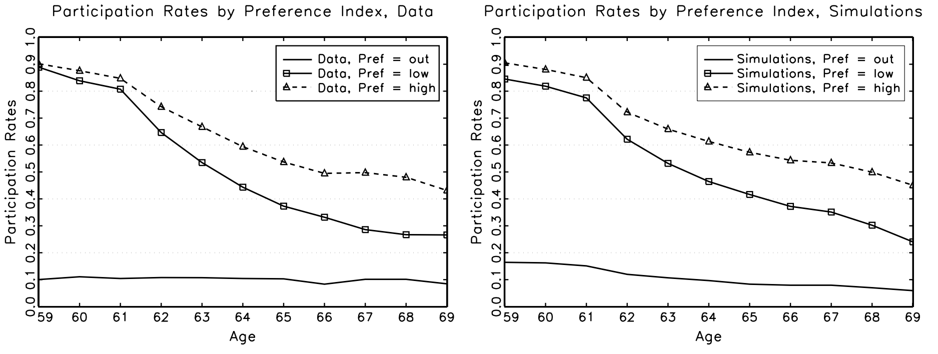

FIGURE 6.—Labor force participation rates by preference index: data and simulations.

This highlights the importance of the preference index in identifying preference heterogeneity.

6.4. *The Effects of Employer-Provided Health Insurance*

The labor supply patterns in Figures 2 and 4 show that those who have retiree coverage retire earlier than those who have tied coverage. However, the profiles do not identify the effects of health insurance on retirement, for three reasons. First, as shown in Table III, those who have retiree coverage have greater pension wealth than other groups. Second, as shown in Figure 1, pension plans for workers who have retiree coverage provide stronger incentives for early retirement than the pension plans held by other groups. Third, as shown in Table V, preferences for leisure vary by health insurance type. In short, retirement incentives differ across health insurance categories for reasons unrelated to health insurance incentives.

To isolate the effects of employer-provided health insurance on labor supply, we conduct some additional simulations. We give everyone the pension accrual rates of tied workers so that pension incentives are identical across health insurance types. We then simulate the model twice, assuming first that all workers have retiree health insurance coverage at age 59 and then assuming tied coverage at age 59. Across the two simulations, households face different medical expense distributions, but in all other dimensions the distribution of incentives and preferences is identical.

This exercise reveals that if all workers had retiree coverage rather than tied coverage, the job exit rate at age 62 would be 8.5 percentage points higher. In contrast, the raw difference in model-predicted exit rates at age 62 is 10.5 percentage points. (The raw difference in the data is 8.2 percentage points.) The high age-62 exit rates of those who have retiree coverage are thus partly due to more generous pensions and stronger preferences for leisure. Even after controlling for these factors, however, health insurance is still an important determinant of retirement.

The effects of health insurance can also be measured by comparing participation rates. We find that the labor force participation rate for ages 60–69 would be 5.1 percentage points lower if everyone had retiree, rather than tied, coverage at age 59. Furthermore, moving everyone from retiree to tied coverage increases the average retirement age (defined as the oldest age at which the individual works plus 1) by 0.34 years.

In comparison, Blau and Gilleskie's (2001) reduced-form estimates imply that having retiree coverage, rather than tied coverage, increases the job exit rate 7.5 percentage points at age 61. Blau and Gilleskie also found that accounting for selection into health insurance plans modestly increases the estimated effect of health insurance on exit rates. Other reduced-form findings in the literature are qualitatively similar to Blau and Gilleskie. For example, Madrian, Burtless, and Gruber (1994) found that retiree coverage reduces the retirement age by 0.4–1.2 years, depending on the specification and the data employed. Karoly and Rogowski (1994), who attempted to account for selection into health insurance plans, found that retiree coverage increases the job exit rate 8 percentage points over a $2\frac{1}{2}$-year period. Our estimates, therefore, lie within the lower bound of the range established by previous reduced-form studies, giving us confidence that the model can be used for policy analysis.

Structural studies that omit medical expense risk find smaller health insurance effects than we do. For example, Gustman and Steinmeier (1994) found that retiree coverage reduces years in the labor force by 0.1 years. Lumsdaine, Stock, and Wise (1994) found even smaller effects. Structural studies that include medical expense risk but omit self-insurance find bigger effects. Our estimated effects are larger than Blau and Gilleskie's (2006, 2008), who found that retiree coverage reduces average labor force participation 1.7 and 1.6 percentage points, respectively, but are smaller than the effects found by Rust and Phelan (1997).[17]

6.5. *Model Validation*

Following several recent studies (e.g., Keane and Wolpin (2007)), we perform an out-of-sample validation exercise. Recall that we estimate the model on a cohort of individuals aged 57–61 in 1992. We test our model by considering the HRS cohort aged 51–55 in 1992; we refer to this group as our validation cohort. These individuals faced different Social Security incentives than did the estimation cohort. The validation cohort did not face the Social Security earnings test after age 65, had a later full retirement age, and faced a benefit

[17] Blau and Gilleskie (2006) considered the retirement decision of couples, and allowed husbands and wives to retire at different dates. Blau and Gilleskie (2008) allowed workers to choose their medical expenses. Because these modifications provide additional mechanisms for smoothing consumption over medical expense shocks, they could reduce the effect of employer-provided health insurance.

TABLE VI

PARTICIPATION RATES BY BIRTH YEAR COHORT

	Data			Model		
	1933	1939	Difference[a]	1933	1939	Difference[a]
60	0.657	0.692	0.035	0.650	0.706	0.056
61	0.636	0.642	0.006	0.622	0.677	0.055
62	0.530	0.545	0.014	0.513	0.570	0.057
63	0.467	0.508	0.041	0.456	0.490	0.035
64	0.408	0.471	0.063	0.413	0.449	0.037
65	0.358	0.424	0.066	0.378	0.459	0.082
66	0.326	0.382	0.057	0.350	0.430	0.080
67	0.314	0.374	0.060	0.339	0.386	0.047
Total, 60–67	3.696	4.037	0.341	3.721	4.168	0.447

[a]The 1939 column minus the 1933 column.

adjustment formula that more strongly encouraged delayed retirement. In addition to facing different Social Security rules, the validation cohort possessed different endowments of wages, wealth, and employer benefits. A useful test of our model, therefore, is to see if it can predict the behavior of the validation cohort.

The Data columns of Table VI show the participation rates observed in the data for each cohort and the difference between them. The data suggest that the change in the Social Security rules coincides with increased labor force participation, especially at later ages. By way of comparison, Song and Manchester (2007), examining Social Security administrative data, found that between 1996 and 2003, participation rates increased by 3, 4, and 6 percentage points for workers turning 62–64, 65, and 66–69, respectively. These differences are similar to the differences between the 1933 and 1939 cohorts in our data, as shown in the fourth column.

The Model columns of Table VI show the participation rates predicted by the model. The simulations for the validation cohort use the initial distribution and Social Security rules for the validation cohort, but use the parameter values estimated on the older estimation cohort.[18] Comparing the diference columns shows that the model-predicted increase in labor supply (0.45 years) resembles the increase observed in the data (0.35 years).

[18]We do not adjust for business cycle conditions. Because the validation cohort starts at age 53, 6 years before the estimation cohort, the validation exercise requires its own wage selection adjustment and pension prediction equation. Using the baseline preference estimates, we construct these inputs in the same way we construct their baseline counterparts. In addition, we adjust the intercept terms in the type prediction equations so that the validation cohort generates the same distribution of preference types as the estimation sample.

7. POLICY EXPERIMENTS

The preceding sections showed that the model fits the data well, given plausible preference parameters. In this section, we use the model to predict how changing the Social Security and Medicare rules would affect retirement behavior. The results of these experiments are summarized in Table VII.

The first data column of Table VII shows model-predicted labor market participation at ages 60–69 under the 1998 Social Security rules. Under the 1998 rules, the average person works a total of 4.29 years over this 10-year period. The last column of Table VII shows that this is close to the total of 4.28 years observed in the data.

The Social Security rules are slowly evolving over time. If current plans continue, by 2030 the normal Social Security retirement age, the date at which workers can receive "full benefits," will have risen from 65 to 67. Raising the normal retirement age to 67 effectively eliminates 2 years of Social Security benefits. The second data column shows the effect of this change.[19] The wealth effect of lower benefits leads years of work to increase by 0.076 years, to 4.37 years.[20]

TABLE VII

EFFECTS OF CHANGING THE SOCIAL SECURITY RETIREMENT AND MEDICARE ELIGIBILITY AGES[a]

	SS = 65 MC = 65	SS = 67[b] MC = 65	SS = 65 MC = 67	SS = 67[b] MC = 67	Data
60	0.650	0.651	0.651	0.652	0.657
61	0.622	0.625	0.623	0.626	0.636
62	0.513	0.526	0.516	0.530	0.530
63	0.456	0.469	0.460	0.472	0.467
64	0.413	0.426	0.422	0.433	0.407
65	0.378	0.386	0.407	0.415	0.358
66	0.350	0.358	0.374	0.381	0.326
67	0.339	0.346	0.341	0.347	0.314
68	0.307	0.311	0.307	0.312	0.304
69	0.264	0.270	0.264	0.270	0.283
Total 60–69	4.292	4.368	4.366	4.438	4.283

[a] SS = Social Security normal retirement age; MC = Medicare eligibility age.
[b] Benefits reduced by 2 years, as described in text.

[19] Under the 2030 rules, an individual claiming benefits at age 65 would receive an annual benefit 13.3% smaller than the benefit he would have received under the 1998 rules (holding AIME constant). We thus implement the 2-year reduction in benefits by reducing annual benefits by 13.3% at every age.

[20] In addition to reducing annual benefits, the intended 2030 rules would impose two other changes. First, the rate at which benefits increase for delaying retirement past the normal age

The third data column of Table VII shows participation when the Medicare eligibility age is increased to 67.[21] Over a 10-year period, total years of work increase by 0.074 years, so that the average probability of employment increases by 0.74 percentage points per year. This amount is larger than the changes found by Blau and Gilleskie (2006), whose simulations show that increasing the Medicare age increases the average probability of employment by 0.1 percentage points, but is smaller than the effects suggested by Rust and Phelan's (1997) analysis.

The fourth data column shows the combined effect of cutting Social Security benefits and raising the Medicare eligibility age. The joint effect is an increase of 0.146 years, 0.072 years more than that generated by cutting Social Security benefits in isolation. In summary, the model predicts that raising the Medicare eligibility age will have almost the same effect on retirement behavior as the benefit reductions associated with a higher Social Security retirement age. Medicare has an even bigger effect on those who have tied coverage at age 59.[22] Simulations reveal that for those who have tied coverage, eliminating 2 years of Social Security benefits increases years in the labor force by 0.12 years, whereas shifting forward the Medicare eligibility age to 67 would increase years in the labor force by 0.28 years.

To understand better the incentives generated by Medicare, we compute the value type-1 individuals place on employer-provided health insurance by finding the increase in assets that would make an uninsured type-1 individual as well off as a person with retiree coverage. In particular, we find the compensating variation $\lambda_t = \lambda(A_t, B_t, H_t, \text{AIME}_t, \omega_t, \zeta_{t-1}, t)$, where

$$\begin{aligned} &V_t(A_t, B_t, H_t, \text{AIME}_t, \omega_t, \zeta_{t-1}, \text{retiree}) \\ &\quad = V_t(A_t + \lambda_t, B_t, H_t, \text{AIME}_t, \omega_t, \zeta_{t-1}, \text{none}). \end{aligned}$$

would increase from 5.5% to 8.0%. This change, like the reduction in annual benefits, should encourage work. However, raising the normal retirement age implies that the relevant earnings test for ages 65–66 would become the stricter, early-retirement test. This change should discourage work. We find that when we switch from the 1998 to the 2030 rules, the effects of the three changes cancel out, so that total hours over ages 60–69 are essentially unchanged.

[21]By shifting forward the Medicare eligibility age to 67, we increase from 65 to 67 the age at which medical expenses can follow the "with Medicare" distribution shown in Table I.

[22]Only 13% of the workers in our sample had tied coverage at age 59. In contrast, Kaiser/HRET (2006) estimated that about 50% of large firms offered tied coverage in the mid-1990s. We might understate the share with tied coverage because, as shown in the Kaiser/HRET study, the fraction of workers with tied (instead of retiree) coverage grew rapidly in the 1990s, and our health insurance measure is based on wave-1 data collected in 1992. In fact, the HRS data indicate that later waves had a higher proportion of individuals with tied coverage than wave 1. We may also be understating the share with tied coverage because of changes in the wording of the HRS questionnaire; see Supplemental Material Appendix H for details.

TABLE VIII

VALUE OF EMPLOYER-PROVIDED HEALTH INSURANCE[a]

	Compensating Assets		Compensating Annuity	
Asset Levels	With Uncertainty	Without Uncertainty	With Uncertainty	Without Uncertainty
Baseline Case				
−$5,700	$20,400	$10,700	$4,630	$2,530
$51,600	$19,200	$10,900	$4,110	$2,700
$147,200	$21,400	$10,600	$4,180	$2,540
$600,000	$16,700	$11,900	$2,970	$2,360
No-Saving Cases[b]				
(a) −$6,000	$112,000	$8,960	$11,220	$2,160
(b) −$6,000	$21,860	$6,860	$3,880	$2,170

[a]Compensating variation between retiree and none coverages for agents with type-1 preferences. Calculations described in text.

[b]No-Saving case (a) uses benchmark preference parameter values; case (b) uses parameter values estimated for no-saving specification.

Table VIII shows the compensating variation $\lambda(A_t, 0, \text{good}, \$32{,}000, 0, 0, 60)$ at several different asset (A_t) levels.[23] The first data column of Table VIII shows the valuations found under the baseline specification. One of the most striking features is that the value of employer-provided health insurance is fairly constant through much of the wealth distribution. Even though richer individuals can better self-insure, they also receive less protection from the government-provided consumption floor. These effects more or less cancel each other out over the asset range of −$5,700 to $147,000. However, individuals with asset levels of $600,000 place less value on retiree coverage, because they can better self-insure against medical expense shocks.

Part of the value of retiree coverage comes from a reduction in average medical expenses—because retiree coverage is subsidized—and part comes from a reduction in the volatility of medical expenses—because it is insurance. To separate the former from the latter, we eliminate medical expense uncertainty, by setting the variance shifter $\sigma(H_t, I_t, t, B_t, P_t)$ to zero, and recompute λ_t, using the same state variables and mean medical expenses as before. Without medical expense uncertainty, λ_t is approximately $11,000. Comparing the two values of λ_t shows that for the typical worker (with $147,000 of assets), about half of the value of health insurance comes from the reduction of average medical expenses and half comes from the reduction of medical expense volatility.

[23]In making these calculations, we remove health-insurance-specific differences in pensions, as described in Section 6.4. It is also worth noting that for the values of H_t and ζ_{t-1} considered here, the conditional differences in expected medical expenses are smaller than the unconditional differences shown in Table I.

The first two data columns of Table VIII measure the lifetime value of health insurance as an asset increment that can be consumed immediately. An alternative approach is to express the value of health insurance as an illiquid annuity comparable to Social Security benefits. The last two columns show this "compensating annuity."[24] When the value of health insurance is expressed as an annuity, the fraction of its value attributable to reduced medical expense volatility falls from 50% (one-half) to about 40%. In most other respects, however, the asset and annuity valuations of health insurance have similar implications.

To summarize, allowing for medical expense uncertainty greatly increases the value of health insurance. It is, therefore, unsurprising that we find larger effects of health insurance on retirement than do Gustman and Steinmeier (1994) and Lumsdaine, Stock, and Wise (1994), who assumed that workers value health insurance at its actuarial cost.

8. ALTERNATIVE SPECIFICATIONS

To consider whether our findings are sensitive to our modelling assumptions, we reestimate the model under three alternate specifications.[25] Table IX shows model-predicted participation rates under the different specifications, along with the data. The parameter estimates behind these simulations are shown in Supplemental Material Appendix K. The first data column of Table IX presents our baseline case. The second column presents the case where individuals are not allowed to save, the third column presents the case with no preference heterogeneity, and the fourth column presents the case where we remove the subjective preference index from the type prediction equations and the GMM criterion function. The last column presents the data. In general, the different specifications match the data profile equally well.

Table X shows how total years of work over ages 60–69 are affected by changes in Social Security and Medicare under each of the alternative specifications. In all specifications, decreasing the Social Security benefits and raising the Medicare eligibility age increase years of work by similar amounts.

[24]To do this, we first find compensating AIME, $\widehat{\lambda}_t$, where

$$\begin{aligned} &V_t(A_t, B_t, H_t, \mathrm{AIME}_t, \omega_t, \zeta_{t-1}, \text{retiree}) \\ &\quad = V_t(A_t, B_t, H_t, \mathrm{AIME}_t + \widehat{\lambda}_t, \omega_t, \zeta_{t-1}, \text{none}). \end{aligned}$$

This change in AIME in turn allows us to calculate the change in expected pension and Social Security benefits that the individual would receive at age 65, the sum of which can be viewed as a compensating annuity. Because these benefits depend on decisions made after age 60, the calculation is only approximate.

[25]In earlier drafts of this paper (French and Jones (2004b, 2007)), we also estimated a specification where housing wealth is illiquid. Although parameter estimates and model fit for this case were somewhat different than our baseline results, the policy simulations were similar.

TABLE IX

MODEL PREDICTED PARTICIPATION BY AGE: ALTERNATIVE SPECIFICATIONS

Age	Baseline	No Saving	Homogeneous Preferences	No Preference Index	Data
60	0.650	0.648	0.621	0.653	0.657
61	0.622	0.632	0.595	0.625	0.636
62	0.513	0.513	0.517	0.516	0.530
63	0.456	0.457	0.453	0.459	0.467
64	0.413	0.429	0.409	0.417	0.407
65	0.378	0.380	0.365	0.381	0.358
66	0.350	0.334	0.351	0.357	0.326
67	0.339	0.327	0.345	0.346	0.314
68	0.307	0.308	0.319	0.314	0.304
69	0.264	0.282	0.286	0.273	0.283
Total 60–69	4.292	4.309	4.260	4.340	4.283

8.1. *No Saving*

We have argued that the ability to self-insure through saving significantly affects the value of employer-provided health insurance. One test of this hypothesis is to modify the model so that individuals cannot save, and examine how labor market decisions change. In particular, we require workers to consume their income net of medical expenses, as in Rust and Phelan (1997) and Blau and Gilleskie (2006, 2008).

The second data column of Table IX contains the labor supply profile generated by the no-saving specification. Comparing this profile to the baseline case shows that, in addition to its obvious failings with respect to asset holdings,

TABLE X

EFFECTS OF CHANGING THE SOCIAL SECURITY RETIREMENT AND MEDICARE ELIGIBILITY AGES, AGES 60–69: ALTERNATIVE SPECIFICATIONS[a]

Rule Specification	Baseline	No Saving	Homogeneous Preferences	No Preference Index
Baseline: SS = 65, MC = 65	4.292	4.309	4.260	4.340
SS = 67: Lower benefits[b]	4.368	4.399	4.335	4.411
SS = 65, MC = 67	4.366	4.384	4.322	4.417
SS = 67[b], MC = 67	4.438	4.456	4.395	4.482

[a] SS = Social Security normal retirement age; MC = Medicare eligibility age.

[b] Benefits reduced by 2 years, as described in text.

the no-saving case matches the labor supply data no better than the baseline case.[26]

Table VIII displays two sets of compensating values for the no-saving case. Case (a), which uses the parameter values from the benchmark case, shows that eliminating the ability to save greatly increases the value of retiree coverage: when assets are −$6,000, the compensating annuity increases from $4,600 in the baseline case (with savings) to $11,200 in the no-savings case (a). When there is no medical expense uncertainty, the comparable figures are $2,530 in the baseline case and $2,160 in the no-savings case. Thus, the ability to self-insure through saving significantly reduces the value of employer-provided health insurance. Case (b) shows that using the parameter values estimated for the no-saving specification, which include a lower value of the risk parameter ν, also lowers the value of insurance.

Simulating the responses to policy changes, we find that raising the Medicare eligibility age to 67 leads to an additional 0.075 years of work, an amount almost identical to that of the baseline specification.

8.2. *No Preference Heterogeneity*

To assess the importance of preference heterogeneity, we estimate and simulate a model where individuals have identical preferences (conditional on age and health status). Comparing the first, third, and last data columns of Table IX shows that the model without preference heterogeneity matches aggregate participation rates as well as the baseline model. However, the no preference heterogeneity specification does much less well in replicating the way in which participation varies across the asset distribution, and, not surprisingly, does not replicate the way in which participation varies across our discretized preference index.

When preferences are homogeneous the simulated response to delaying the Medicare eligibility age, 0.062 years, is similar to the response in the baseline specification. This is consistent with our analysis in Section 6.4, where not accounting for preference heterogeneity and insurance self-selection appeared to only modestly change the estimated effects of health insurance on retirement.

8.3. *No Preference Index*

In the baseline specification, we use the preference index (described in Section 4.4) to predict preference type, and the GMM criterion function includes participation rates for each value of the index. Because labor force participation differs sharply across the index in ways not predicted by the model's other

[26]Because the baseline and no-saving cases are estimated with different moments, their overidentification statistics are not comparable. However, inserting the decision profiles generated by the baseline model into the moment conditions used to estimate the no-saving case produces an overidentification statistic of 354, while the no-saving specification produces an overidentification statistic of 366.

state variables, we interpret the index as a measure of otherwise unobserved preferences toward work. It is possible, however, that using the preference index causes us to overstate the correlation between health insurance and tastes for leisure. For example, Table III shows that employed individuals with retiree coverage are more likely to have a preference index that is low than employed individuals with tied coverage. This means that workers with retiree coverage are more likely to report looking forward to retirement, and thus more likely to be assigned a higher desire for leisure. But workers with retiree coverage may be more likely to report looking forward to retirement simply because they would have health insurance and other financial resources during retirement. As a robustness test, we remove the preference index and the preference index-related moment conditions, and reestimate the model.

Table XI contains summary statistics for the preference groups generated by this alternative specification. Comparing Table XI to the baseline results contained in Table V reveals that eliminating the preference index from the type prediction equations changes only modestly the parameter estimates and the distribution of insurance coverage across the three preference types. The model without the preference index provides less evidence of self-selection: when the preference index is removed the fraction of high preference for work, type-2 individuals with tied coverage falls from 15.8% to 8.9%.

Table X shows that excluding the preference index only slightly changes the estimated effect of Medicare and Social Security on labor supply. Given that self-selection has only a small effect on our results when we include the pref-

TABLE XI

MEAN VALUES BY PREFERENCE TYPE: ALTERNATIVE SPECIFICATION

	Type 0	Type 1	Type 2
Key Preference Parameters			
γ	0.405	0.647	0.986
β	0.962	0.858	1.143
Means by Preference Type			
Assets ($1,000s)	115	231	376
Pension Wealth ($1,000s)	60	108	85
Wages ($/hour)	11.0	18.4	13.5
Probability of Health Insurance Type, Given Preference Type			
Health insurance = none	0.392	0.193	0.394
Health insurance = retiree	0.560	0.633	0.518
Health insurance = tied	0.047	0.174	0.089
Probability of Preference Index Value, Given Preference Type			
Preference index = out	0.523	0.216	0.224
Preference index = low	0.247	0.399	0.363
Preference index = high	0.230	0.385	0.413
Fraction with preference type	0.246	0.635	0.119

erence index, it should come as no surprise that self-selection has only a small effect when we exclude the index.

9. CONCLUSION

Prior to age 65, many individuals receive health insurance only if they continue to work. At age 65, however, Medicare provides health insurance to almost everyone. Therefore, a potentially important work incentive disappears at age 65. To see if Medicare benefits have a large effect on retirement behavior, we construct a retirement model that includes health insurance, uncertain medical costs, a savings decision, a nonnegativity constraint on assets, and a government-provided consumption floor.

Using data from the Health and Retirement Study, we estimate the structural parameters of our model. The model fits the data well, with reasonable preference parameters. In addition, the model does a satisfactory job of predicting the behavior of individuals who, by belonging to a younger cohort, face different Social Security rules than the individuals on which the model was estimated.

We find that health care uncertainty significantly affects the value of employer-provided health insurance. Our calculations suggest that about half of the value workers place on employer-provided health insurance comes from its ability to reduce medical expense risk. Furthermore, we find evidence that individuals with higher tastes for leisure are more likely to choose employers who provide health insurance to early retirees. Nevertheless, we find that Medicare is important for understanding retirement, especially for workers whose health insurance is tied to their job. For example, the effects of raising the Medicare eligibility age to 67 are just as large as the effects of reducing Social Security benefits.

APPENDIX A: CAST OF CHARACTERS

Preference Parameters		Health-Related Parameters	
γ	Consumption weight	H_t	Health status
β	Time discount factor	M_t	Out-of-pocket medical expenses
ν	Coefficient of relative risk aversion, utility	I_t	Health insurance type
		$m(\cdot)$	Mean shifter, logged medical expenses
θ_B	Bequest weight	$\sigma(\cdot)$	Volatility shifter, logged medical expenses
κ	Bequest shifter	ψ_t	Idiosyncratic medical expense shock
$C_{\min}$	Consumption floor	ζ_t	Persistent medical expense shock
L	Leisure endowment	ε_t	Innovation, persistent shock
ϕ_H	Leisure cost of bad health	ρ_m	Autocorrelation, persistent shock
ϕ_{Pt}	Fixed cost of work	σ_ε^2	Innovation variance, persistent shock
ϕ_{P0}	Fixed cost, intercept	ξ_t	Transitory medical expense shock
ϕ_{P1}	Fixed cost, time trend	σ_ξ^2	Variance, transitory shock
ϕ_{RE}	Reentry cost		

Decision Variables		Wage-Related Parameters	
C_t	Consumption	W_t	Hourly wage
N_t	Hours of work	$W(\cdot)$	Mean shifter, logged wages
L_t	Leisure	α	Coefficient on hours, logged wages
P_t	Participation	ω_t	Idiosyncratic wage shock
A_t	Assets	ρ_W	Autocorrelation, wage shock
B_t	Social Security application	η_t	Innovation, wage shock
		σ_η^2	Innovation variance, wage shock
Financial Variables			
$Y(\cdot)$	After-tax income	Miscellaneous	
τ	Tax parameter vector	s_t	Survival probability
r	Real interest rate	pref	Discrete preference index
ys_t	Spousal income	X_t	State vector, worker's problem
$ys(\cdot)$	Mean shifter, spousal income	$\lambda(\cdot)$	Compensating variation
ss_t	Social Security income	T	Number of years in GMM criterion
AIME_t	Social Security wealth		
pb_t	Pension benefits		

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