What Accounts for the Variation in Retirement Wealth Among U.S. Households?

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Even among households with similar socioeconomic characteristics, saving and wealth vary considerably. Life-cycle models attribute this variation to differences in time preference rates, risk tolerance, exposure to uncertainty, relative tastes for work and leisure at advanced ages, and income replacement rates. These factors have testable implications concerning the relation between accumulated wealth and the shape of the consumption profile. Using the Panel Study of Income Dynamics and the Consumer Expenditure Survey, we find little support for these implications. The data are instead consistent with "rule of thumb," "mental accounting," or hyperbolic discounting theories of wealth accumulation. (*JEL* D1, D91, E21)

Wealth and saving vary considerably, even among households with similar socioeconomic characteristics (Steven Venti and David Wise, 1998; Annamaria Lusardi, 1999). The interpretation of this variation is a pivotal issue. If saving reflects rational, farsighted optimization, then low saving is simply an expression of preferences—saying that someone saves "too little" is comparable to asserting that he or she doesn't listen to enough classical music (Edward Lazear, 1994). If, however, households are shortsighted, boundedly rational, dynamically inconsistent, impulsive, or prone to regret, then the adequacy of saving is a well-posed and important empirical issue (B. Douglas Bernheim, 1995).

Various factors could in principle account for the observed variation in savings for retirement within the context of standard life-cycle models with rational, farsighted optimization. Households may differ with respect to patience (the rate of pure time preference), risk tolerance, exposure to uncertainty, health status, perceived life expectancy, relative tastes for goods complementary with leisure at advanced ages, levels of work-related expenses, lifetime earnings, or income replacement rates. In this paper, we test for the presence of these factors by studying data on wealth, income, and consumption drawn from the Panel Study of Income Dynamics (PSID) and the Consumer Expenditure Survey (CEX).

Explanations for the variation in wealth that are based on the life-cycle framework fall into three broad categories. Each category has a distinctive and testable empirical implication. Factors in the first category create systematic correlations between the household's wealth and its average *consumption growth rate.* For example, if differences in saving result from differences in patience, those who save more should also exhibit higher consumption growth rates. Factors in the second category create systematic correlations between wealth and *one-time changes in consumption at retirement.* For example, households with higher work-related expenses should accumulate less wealth and experience larger declines in measured consumption at retirement. Likewise, an unexpected early retirement prematurely terminates

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wealth accumulation, and should depress consumption by forcing the retiree to revise his or her expectations about lifetime resources. Factors in the third category give rise to systematic correlations between accumulated wealth and the *level of consumption.* Households with strong bequest motives, for example, should tend to consume less throughout the life cycle.

Our first central finding is that there is essentially no relation between accumulated wealth and consumption growth rates either prior to retirement or after retirement. This suggests that a wide range of factors (including differences in patience, as measured by pure rates of time preference) fail to provide even contributory explanations for the observed variations in accumulated wealth.

Our second central finding is the existence of a correlation between wealth and the decline in consumption at retirement. This is superficially consistent with the second category of explanations mentioned previously. However, these hypotheses do not survive closer scrutiny. The impact of retirement on work-related expenses and leisure substitutes is too small to explain the decline in consumption, and other expenditures also fall sharply. Moreover, there is little evidence that the decline in potentially work-related expenses is larger for households with less wealth. Nor is our second finding the consequence of unanticipated events that affect the timing of retirement: the pattern persists even when we remove these effects statistically. Taken together, our first two findings imply that a broad range of standard life-cycle considerations are collectively incapable of accounting for the observed variation in wealth, holding constant income profiles.

Differences in income profiles should also contribute to variation in wealth. For instance, households with relatively generous pensions and social security benefits should accumulate less wealth to smooth consumption through retirement. Yet in practice, we find that households with lower income replacement rates do not have significantly higher wealth/income ratios. One can explain this finding in the context of life cycle theory if pension coverage is correlated with tastes for saving. However, it is much more difficult to account for the observed pattern in light of our third central finding: the discontinuous drop in consumption at retirement is larger for households with relatively less generous pension and social security benefits. Although our three central results challenge the validity of standard life-cycle models, they do not rule out a variety of behavioral theories. We discuss these in more detail below.

This paper is related to several previous studies. Daniel Hamermesh (1984) and Randall Mariger (1987) find that consumption declines sharply as households move into retirement. Hamermesh also infers that consumption at retirement is not sustainable; Laurence Kotlikoff et al. (1982) dispute this conclusion. A. Leslie Robb and John Burbridge (1989) find that, among Canadians, consumption at retirement falls more sharply for blue-collar workers than for white-collar workers. Jerry Hausman and Lynn Paquette (1987) link the decline in (non-medical) consumption at retirement to unexpected and involuntary job loss, often resulting from health problems. James Banks et al. (1998) track consumption and earnings of synthetic British cohorts through retirement years and document a sharp drop in average consumption at age 65. They argue that this finding is difficult to explain with reference to conventional economic factors such as a reduction in anticipated labor force participation or work-related expenses. Eric Engen et al. (1999) maintain that it is possible to account for the observed variation in wealth using a life-cycle model with heterogeneous earnings chocks and pension coverage. However, their model presupposes consumption smoothing at retirement, which is inconsistent with the behavioral patterns documented below.

I. Theoretical Preliminaries

In discussing the potential sources of variation in retirement wealth, it is useful to distinguish between variation in income profiles and all other factors. We consider these in reverse order.

A. *Sources of Variation in Wealth, Fixing Income Profiles*

For households with similar earnings histories, pensions, inheritances, and retirement ages, those with less wealth at retirement consumed more prior to retirement, and will consume or bequeath less after retirement. By itself, the budget constraint does not tie down the characteristics of the consumption profile more precisely. For specific models, the consumption profile accommodates the budget constraint in one (or more) of three ways. First, for those with lower wealth at retirement, consumption may grow less rapidly over the life cycle. Second, consumption may decline discontinuously at retirement, and this discontinuity may be larger for those with less accumulated wealth. Third, those with less accumulated wealth at retirement may bequeath less, consuming more throughout their lives.[1] By studying the relations between accumulated wealth and consumption profiles, one can therefore identify the factors that do and do not contribute to the observed variation in wealth.

1. *Factors Affecting the Slope of the Consumption Profile.*—Variation in accumulated wealth at retirement could in principle result from any factor that produces variation in the slope of the consumption profile. Subject to the qualifications discussed below, rising consumption profiles correspond to high accumulation, whereas falling consumption profiles correspond to low accumulation.

To illustrate, consider a time-separable utility function of the form

$$(1)\quad U_t = U(C_t) + E_t\left\{ \sum_{s=t+1}^{\infty} \lambda_{t,s} U(C_s) \right\},$$

with

$$(2)\quad \lambda_{t,s} = \left(\frac{1}{1+\delta}\right)^{s-t} \prod_{k=t+1}^{s} (1-\pi_k),$$

where C_i is consumption in period i, E_t is the expectations operator (conditional on information available at time t), δ is the standard (constant) pure rate of time preference, and π_k is the probability of dying before period k, conditional upon surviving to period $k-1$. Assume for the moment that income is potentially uncertain and independently distributed across periods, but that the age of retirement is fixed.

Maximization of (1) subject to a resource constraint yields the following first-order condition:

$$(3)\quad U'(C_t) = \alpha_t E_t\{U'(C_{t+1})\},$$

where

$$(4)\quad \alpha_t \equiv \frac{(1+r)(1-\pi_{t+1})}{1+\delta}.$$

Taking a second-order Taylor series expansion of (3) yields the familiar Euler equation

$$(5)\quad \frac{E_t(C_{t+1}) - C_t}{C_t} \approx \gamma\left(1 - \frac{1}{\alpha_t}\right) + \frac{\psi}{2}\sigma_{t+1}^2,$$

where γ is the intertemporal elasticity of substitution, ψ reflects the household's precautionary inclinations, and σ_{t+1} is the standard deviation of consumption in period $t+1$. Equation (5) tells us that the slope of the consumption profile depends on the rate of interest, the pure rate of time preference (patience), perceived survival probabilities, and a precautionary motive, as captured in the final term (see, e.g., Angus Deaton, 1991; Karen Dynan, 1993; Michael Palumbo, 1999). Variation in any of these factors can produce variation in the shape of the consumption profile and associated variation in wealth.

Consider the effects of variation in patience. If the elasticity of substitution is close to zero, then variation in δ does not alter the consumption trajectory, and therefore cannot account for observed variation in wealth. Early studies placed γ near zero (Robert Hall, 1988), but more recent evidence supports a small positive elasticity (Orazio Attanasio and Guglielmo Weber, 1993, 1995). Consequently, one would ex-

[1] As a matter of logic, there are obviously other possibilities; for example, consumption may decline discontinuously at some point after retirement, and this point may occur at a more advanced age for those who reach retirement with greater wealth. This pattern might arise in a model with heterogeneity in finite, deterministic life spans. The assumption of a deterministic life span is, however, unattractive. More generally, variation in survival probabilities gives rise to variations in the slope of the consumption profile—an example of the first pattern.

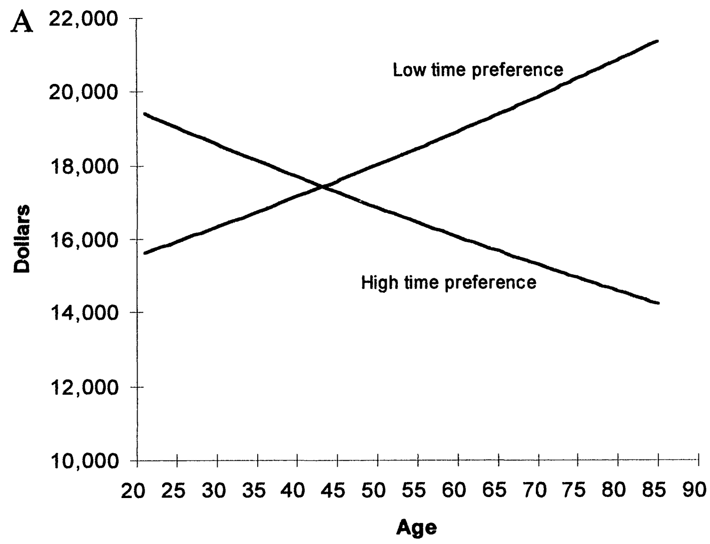

FIGURE 1A. AGE–CONSUMPTION PROFILES FOR HIGH- AND LOW-TIME-PREFERENCE HOUSEHOLDS

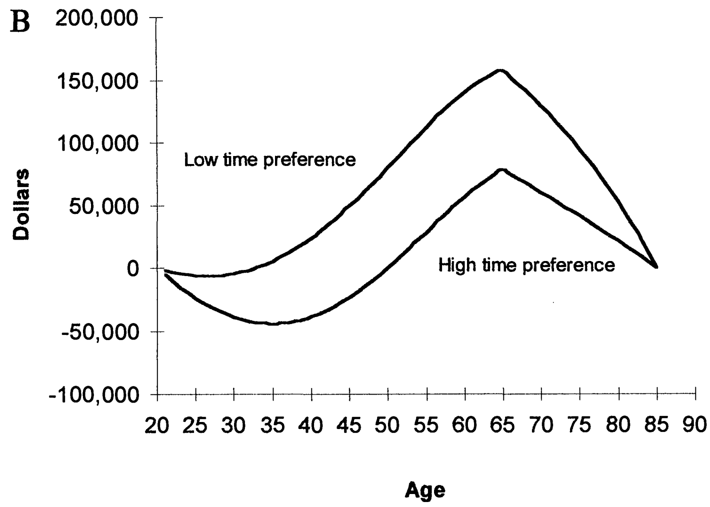

FIGURE 1B. AGE–WEALTH PROFILES FOR HIGH- AND LOW-TIME-PREFERENCE HOUSEHOLDS

pect the slope of the consumption profile to vary with δ. All else equal, impatient households should consume more than patient households early in life, and less later in life. Patient households should therefore accumulate more wealth for retirement (see Figure 1). This produces a positive correlation between the slope of the consumption profile and accumulated wealth at retirement.

For a liquidity-constrained household or a "buffer-stock" saver, the income growth rate dictates the consumption growth rate (Deaton, 1991; Chris Carroll, 1997). Thus, consideration of liquidity constraints disrupts the clean prediction of the basic model (a positive correlation between consumption growth and assets accumulated at retirement). However, because households approaching retirement generally hold nontrivial assets and anticipate significant near-term declines in nonasset income (Banks et al., 1998; Attanasio et al., 1999), we doubt that this is a significant problem for us in practice, except possibly at the lowest end of the wealth distribution.

Equation (5) implies that, like variations in the pure rate of time discount, variations in perceived longevity (survival probabilities), income uncertainty, and tastes for precaution rotate the consumption profile.[2] Thus, variations in wealth accumulation that result from any of these factors should manifest themselves through a positive correlation between retirement wealth and the growth rate of consumption either before retirement, after retirement, or both. The absence of such correlations would be inconsistent with the hypothesis that these factors contribute to the observed variation in wealth.

So far, we have focused exclusively on the standard life-cycle framework. Additional considerations emerge in alternative behavioral models (e.g., Richard Thaler and Hersh Shefrin, 1981; Shefrin and Thaler, 1988; George Akerlof, 1991; David Laibson, 1997). For example, following Laibson, one can introduce dynamic inconsistency by inserting a parameter, $\beta < 1$, measuring the extent to which individuals discount all future utility relative to today's utility, into equation (1):

$$(6) \quad U_t = U(C_t) + \beta \left\{ \sum_{s=t+1}^{\infty} \lambda_{t,s} U(C_s) \right\}.$$

Behavior then corresponds to the equilibrium of a game played by successive incarnations of the decision maker. As Laibson shows, there is a tendency for dynamically inconsistent planners to save too little, exhibiting apparently high rates of time preference. If differences in time consistency explain the observed variations in retirement wealth, one would therefore still expect to find a positive correlation between retirement wealth and the consumption growth rate. However, it is important to emphasize that the implications of these models for retirement saving are still imperfectly understood (Laibson and Christopher Harris, 1998; Bernheim et al., 1999).

2. *Factors Affecting the Change in Consumption at Retirement.*—Variation in wealth may also be attributable to factors that produce downward, discontinuous jumps in consumption at retirement. By the logic of the budget constraint, the existence of this discontinuity generally implies higher consumption, and therefore less wealth accumulation, before retirement. Figure 2 illustrates. We consider two households that are identical in all respects, except that one prefers a flat consumption profile, whereas the other prefers a consumption profile with a consumption discontinuity at retirement. Note that the latter household holds less wealth throughout the life cycle. If such factors account for the observed variation in wealth, one should observe a negative correlation between retirement wealth and the size of the discontinuity.

Factors that can generate sharp changes in consumption at retirement include the existence of work-related expenses and/or preferences for leisure substitutes or complements (William Ghez and Gary Becker, 1975; Banks et al., 1998; Marianne Baxter and Urban Jermann, 1999; and Monika Butler, 2001). Variation in these factors may therefore provide a contributory explanation for differences in retirement wealth. If this explanation is valid, the drop in spending at retirement should be larger for work-related expenditures than that for non-work-related expenditures, and larger for leisure substitutes than for leisure complements. One should also observe a stronger negative correlation between retirement wealth and the absolute size of the discontinuity for expenditure categories that are more closely work-related and more substitutable for leisure. In the absence of these patterns, we would infer that this explanation does account for a significant fraction of the variation in retirement wealth.

Variation in retirement wealth may also result from unexpected events that affect the timing of retirement, such as sudden deterioration of health or loss of job (Peter Diamond and Hausman, 1984; Hausman and Paquette, 1987). For those who retire unexpectedly early—and as a consequence retire with less than desired wealth—retirement coincides with "news" of a negative shock to permanent income. This gives

[2] These factors may change over the life cycle. For example, households may have similar survival probabilities and consumption growth rates but different levels of wealth today because they expect to have different survival probabilities (and hence different consumption growth rates) in the future.

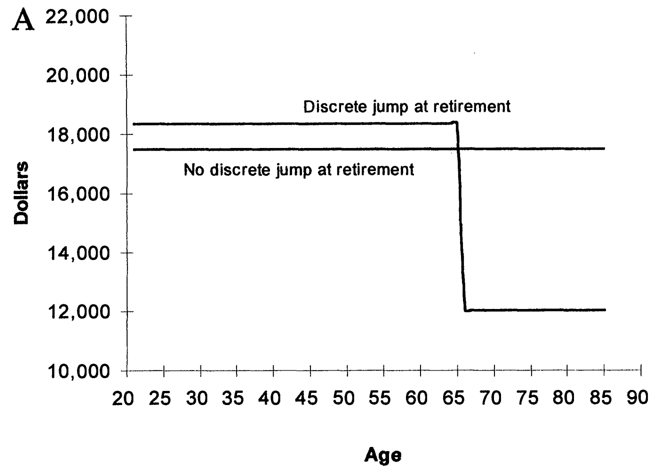

FIGURE 2A. AGE–CONSUMPTION PROFILE WITH (AND WITHOUT) A DISCRETE SHIFT IN CONSUMPTION AT RETIREMENT

FIGURE 2B. AGE–WEALTH PROFILE WITH (AND WITHOUT) A DISCRETE SHIFT IN CONSUMPTION AT RETIREMENT

rise to a negative correlation between retirement wealth and the decline in consumption at retirement. One can test this hypothesis by studying residual changes in retirement consumption after controlling for unexpected retirement shocks.[3] We elaborate on this point in Section II.

Stepping outside the life-cycle framework, one can imagine other explanations for variations in retirement wealth that are related to differences in the size of the consumption discontinuity at retirement. Suppose, for example, that saving is somewhat haphazard, but that individuals evaluate their finances upon reaching retirement, adjusting spending as necessary to accommodate resources. In that case, the adequacy of savings at retirement is "news."

[3] Banks et al. (1998) examine changes in consumption around age 65 for British workers. Because the retirement hazard spikes at this age, their procedure removes the idiosyncratic news from retirement events, and implicitly yields an estimate of the correlation between a predictable determinant of retirement (being 65) and changes in consumption. In the same spirit, we implement a two-stage procedure that allows us to estimate the relation between consumption and predictable changes in retirement.

Those with bad news (inadequate savings) presumably decrease their consumption, whereas those with good news (excessive savings) increase it. Alternatively, individuals may regard current income as more "spendable" than assets, particularly those held in retirement accounts or converted into annuities (Thaler and Shefrin, 1981; Shefrin and Thaler, 1988; Thaler, 1994). Consumption might then decline at retirement simply because current income falls. If individuals differ in the extent to which they exercise self-discipline over spendable funds, then those with little self-discipline should accumulate less wealth for retirement and experience greater declines in spending than those with greater self-discipline. For both of these hypotheses, the predicted patterns (a drop in spending at retirement and a negative correlation between wealth and the size of this discontinuity) should cut across consumption categories, and should emerge even when retirement is predictable. Thus, in principle, it is possible to distinguish these alternatives empirically from the standard life cycle factors discussed previously.[4]

3. *Factors Affecting the Overall Level of Consumption.*—Variation in retirement wealth could result from factors that influence the overall level of consumption throughout the life cycle, with bequests adjusting to satisfy the budget constraint (Dynan et al., 2000). Figure 3 illustrates hypothetical consumption and wealth profiles for an individual with no bequest motive, and one who wishes to bequeath some portion of lifetime wealth. At every age, wealth is higher for the second household than for the first. Thus variation in wealth might occur if, for example, there are significant differences in the strength of bequest motives across the population [for supporting empirical evidence see Bernheim (1991) and John Laitner and F. Thomas Juster (1996)]. Note that theories of this kind have no implications concerning the shape of the consumption profile, or its relation to wealth.

B. *Variation in Wealth Resulting from Differences in Income Profiles*

Survey data generally show that wealth at any age rises significantly with proxies for lifetime resources, such as household earnings. Although this pattern certainly helps to account for the fact that some households accumulate more wealth than others, it fails to discriminate between interesting behavioral hypotheses. Variations in retirement wealth may also result from variations in the *shape* of the income profile. This observation motivates a more promising line of inquiry.

One important characteristic of the income profile is the ratio of preretirement to postretirement nonasset income (the earnings replacement rate). In practice, social security and private pension benefits account for the bulk of postretirement nonasset income, and information on the size of these benefits is readily available. Consequently, households nearing retirement should be able to predict earnings replacement rates to a high degree of accuracy.

Standard life-cycle models imply that, fixing preferences, household saving should vary inversely with predictable differences in earnings replacement rates. In practice, these rates are probably correlated with unobserved household characteristics, such as tastes for saving [pensions are more common among the thrifty (William G. Gale, 1998)]. However, regardless of earnings replacement rates or tastes for saving, the life cycle model implies that households should still adjust saving to smooth consumption. Consequently, even if earnings replacement rates are related to tastes for saving, they

[4] We comment briefly on three other potential sources of correlation between wealth and predictable consumption changes at retirement. First, R. Glenn Hubbard et al. (1995) suggest that asset-based means testing for welfare programs could induce lower income households to hold very low (or zero) levels of wealth. Their model predicts a sharp consumption decline at retirement, but only for the lowest income group; households in the upper half of the income distribution smooth consumption through retirement. We nevertheless find evidence of a negative correlation between wealth and the decline in consumption at retirement even among households in the top half of the income distribution. Second, buffer-stock models imply that consumption may track predictable components of income (Deaton, 1991; Carroll, 1997). However, these models typically predict positive savings and consumption smoothing around retirement. Third, different households may anticipate different changes in relative prices after retirement, resulting perhaps from discount programs that target the elderly. We doubt these discounts are large enough to account for the observed changes in spending, particularly in expenditure categories such as groceries (see the discussion of food consumption, below). Also, eligibility for discounts is typically tied to age, rather than work status.

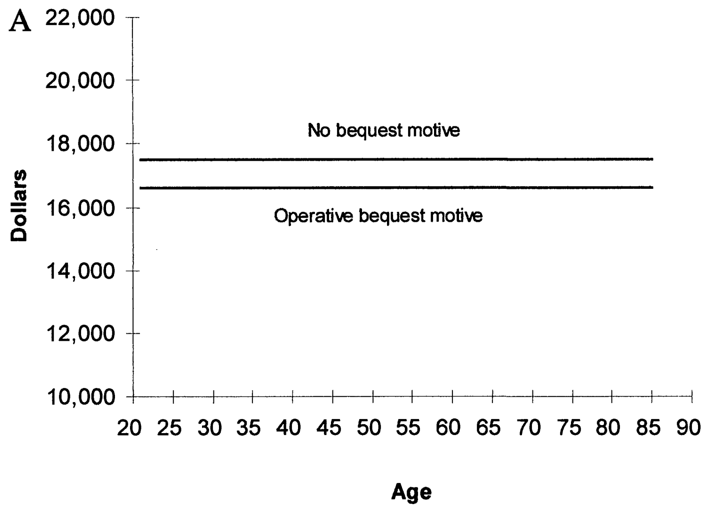

FIGURE 3A. AGE–CONSUMPTION PROFILES WITH (AND WITHOUT) A BEQUEST MOTIVE

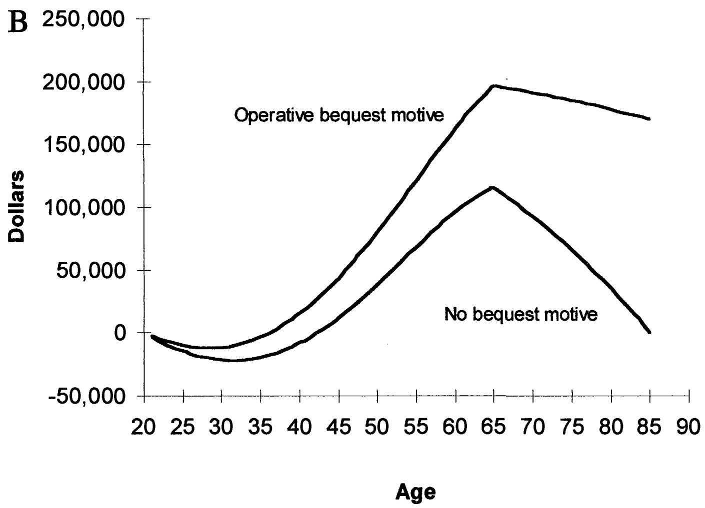

FIGURE 3B. AGE–WEALTH PROFILES WITH (AND WITHOUT) A BEQUEST MOTIVE

should not be correlated with the size of the decline in consumption at retirement, at least among those with nonnegligible retirement savings.

This conclusion must be modified if tastes for thrift are also related to the magnitude of the consumption discontinuity at retirement. We have already discussed the factors that could generate such a relation (Section I, subsection A.2). A spurious correlation between earnings replacement rates and the size of the consumption discontinuity at retirement could arise if workers with tastes for leisure substitutes self-select into jobs with poor pension coverage, or if employers offer less generous pensions to workers with higher work-related expenses. As before, one can test these hypotheses by examining consumption profiles for disaggregated expenditure categories.

II. Empirical Models and Estimation Strategy

As explained in Section I, one can distinguish between competing explanations for the observed variation in wealth accumulation by comparing aspects of consumption dynamics (growth rates and discontinuities at retirement) across households with differing levels of assets and rates of earnings replacement. In this section, we outline our empirical strategy for making these comparisons.

Our general approach is to estimate functions of the following form:

$$(7)\quad \Delta \ln(C_{it}) = \xi(t, \mathbf{X_i}) + \mathbf{\Gamma \Delta Z_{it}} + \zeta_{it}.$$

In this equation, $\Delta y_t \equiv y_{t+1} - y_t$ for any variable y, time t is measured relative to retirement ($t = 0$ corresponds to the year of retirement), C_{it} represents the level of consumption in period t, ξ is a function (discussed in greater detail below), $\mathbf{X_i}$ is a vector of fixed household characteristics, $\mathbf{\Gamma}$ is a vector of parameters, $\mathbf{Z_{it}}$ is a vector of demographic factors that may change through time (such as the number of individuals in the household), and ζ_{it} is a disturbance term.

It is natural to think of (7) as a consumption Euler equation, augmented to allow for shifts in family composition and other demographic determinants of marginal utility (Attanasio et al., 1999). Indeed, it is possible to derive this expression from a simple model in which single-period utility is given by $U(C_{it}\exp(\mathbf{\Gamma Z_{it}}))$ and $U(\cdot)$ belongs to the CRRA family of utility functions, or more generally as an approximation of the Euler equation. Within this familiar framework, the function $\xi(t, \mathbf{X_i})$ captures the effects of household preferences (such as the pure rate of time preference), environmental parameters (such as life expectancy, income uncertainty, and rates of return), and aging (e.g., through its effects on survival probabilities) on the slope of the consumption profile.

If one adopts the Euler equation interpretation of (7), then it is natural to assume that the consumption shock ζ_{it} is serially uncorrelated and independent of all information available to the household at time t. However, if household consumption is subject to measurement error, the estimated residuals for equation (7) may exhibit negative serial correlation. Consequently, when we estimate equation (7), we use clustered Huber–White standard errors (Jeffrey Wooldridge, 2001, Chapter 10) to account for the correlated structure of the within-household covariance matrix.

Our primary objective is to estimate the function $\xi(t, \mathbf{X_i})$. This function describes the average consumption growth rate for a household with characteristics $\mathbf{X_i}$ at age t, removing the effects of any changes in the household's demographic characteristics (that is, assuming $\mathbf{\Delta Z_{it}} = 0$). In this sense, we think of $\xi(t, \mathbf{X_i})$ as household i's "baseline" consumption growth rate at age t. By estimating a sufficiently flexible functional form for ξ, we can determine whether consumption growth rates are systematically correlated with wealth accumulation, and we can ascertain the extent to which consumption discontinuities at retirement are correlated with assets and income replacement ratios.

Our specific approach is to assume that

$$(8)\qquad \xi(t, \mathbf{X_i}) = \mathbf{X_i}(\boldsymbol{\beta}^{\mathbf{W}} I(t < 0) + \boldsymbol{\beta}^{\mathbf{0}} I(t = 0) + \boldsymbol{\beta}^{\mathbf{1}} I(t = 1) + \boldsymbol{\beta}^{\mathbf{R}} I(t > 1)),$$

where $I(\cdot)$ is an indicator function that returns a value of unity when the expression is satisfied and zero otherwise, and $\boldsymbol{\beta}^{\mathbf{W}}$, $\boldsymbol{\beta}^{\mathbf{0}}$, $\boldsymbol{\beta}^{\mathbf{1}}$, and $\boldsymbol{\beta}^{\mathbf{R}}$ are vectors of parameters. In effect, we allow households to have four different consumption growth rates pertaining (respectively) to the years before retirement, the first year of retirement, the second year of retirement, and subsequent years. Moreover, each of these four growth rates varies with household characteristics. We estimate separate growth rates for the first and second years of retirement to allow for the possibility that some adjustments to changes in work status may be delayed.

Though equation (7) is our primary empirical specification, in some instances we also estimate functions of the following form:

$$(9)\quad \ln(C_{it}) = \mu_i + \psi(t, \mathbf{X_i}) + \mathbf{\Omega Z_{it}} + \eta_{it},$$

where μ_i is a household fixed effect, ψ is a function (discussed in greater detail below), $\mathbf{\Omega}$ is a vector of parameters, η_{it} is a mean-zero random variable, and all other symbols are defined as before. It is important to understand that equations (7) and (9) are mutually consistent. To see this, define τ and T, respectively, as the first and last year in which the household is observed. Let $v_{it} \equiv \sum_{k=\tau+1}^{t} \zeta_{it}$ (the cumulative innovation in consumption since time τ) and $\bar{v}_i \equiv (T - \tau)^{-1} \sum_{k=\tau+1}^{T} v_{ik}$ (household i's average consumption innovation). Then (9) follows directly from (7), with $\mu_i \equiv \ln(C_{i\tau}) + \bar{v}_i$, $\eta_{it} \equiv v_{it} - \bar{v}_i$, and $\psi(t, \mathbf{X_i}) \equiv \sum_{k=\tau+1}^{t} \xi(k, \mathbf{X_i})$. Notice that the individual fixed effect μ_i reflects both household specific permanent income [as measured by $\ln(C_{i\tau})$] as well as the household's average consumption innovation $\bar{v}_i$. Likewise, in constructing the error term η_{it}, we subtract $\bar{v}_i$. These operations are necessary because, in a limited number of years, the realized average of the error terms v_{it} is not zero for each household; particular households may on average experience positive or negative innovations. Removing $\bar{v}_i$ from the error term assures us that η_{it} has a mean of zero.[5]

When we estimate equation (9) rather than equation (7), our primary interest is in the function ψ (specifically, the manner in which this function varies with household characteristics). Accordingly, we adopt the following flexible functional form:

$$\psi(t, \mathbf{X_i}) = \mathbf{X_i}\boldsymbol{\beta}_\mathbf{t}, \tag{10}$$

where $\boldsymbol{\beta}_\mathbf{t}$ is a vector of parameters. Note that this formulation permits the parameter vector to vary freely with age relative to retirement t. Thus, estimates of ψ imply an expected consumption trajectory for each household that depends on the fixed characteristics of the household. Even though it is possible in principle to estimate a separate $\boldsymbol{\beta}_\mathbf{t}$ each value of t, in practice this requires the estimation of a very large number of parameters. We therefore impose the restriction that $\boldsymbol{\beta}_\mathbf{t}$ is constant within consecutive two-year intervals. Note that equation (9), unlike equation (7), permits us to use every observation for each household, including data from nonconsecutive years, even when information is unavailable for the intervening year(s).

In all likelihood, the error terms in equation (9), η_{it}, are heteroskedastic and autocorrelated. This is certainly the case if the ζ_{it} are independent, as implied by the permanent income hypothesis. As before, clustered robust Huber–White standard errors are used to correct for the resulting nondiagonal heteroscedastic covariance matrix.

To this point, we have assumed that $\mathbf{X_i}$, $\mathbf{Z_{it}}$, and the timing of retirement are exogenous with respect to the consumption innovations. These assumptions are potentially problematic. To the extent $\mathbf{X_i}$ measures characteristics at retirement, it may be correlated with consumption shocks occurring before retirement. Furthermore, unanticipated changes in $\mathbf{Z_{it}}$ and in retirement plans may be related to permanent per capita income shocks, which affect consumption.[6] We discuss each of these possibilities.

The potential endogeneity of $\mathbf{Z_{it}}$ is least troublesome. If unanticipated changes in these variables (e.g., the death of a family member) are indeed related to permanent per capita income shocks, then one simply reinterprets the coefficient vector $\boldsymbol{\Gamma}$ (or $\boldsymbol{\Omega}$) as reflecting, at least in part, the impact of demographic "news" on consumption. Because we are only concerned with the properties of the function ξ (or ψ), this reinterpretation of $\boldsymbol{\Gamma}$ (or $\boldsymbol{\Omega}$) is innocuous for our purposes.

The potential endogeneity of $\mathbf{X_i}$ is a more serious concern. To test the hypotheses outlined in Section II, we choose the elements of $\mathbf{X_i}$ to include measures of wealth and income replacement rates at retirement, which may be related to consumption shocks that occurred prior to retirement. For the reasons discussed in Section I, subsection B, we doubt that these effects are quantitatively important for income replacement rates. However, it is possible that wealth at retirement and consumption shocks before retirement are related to common third factors. The primary candidates for these third factors

[5] Normally, one should be cautious about estimating random-walk models with fixed effects. The problem is not serious in this case, however, because our primary objective is simply to quantify average consumption relative to a benchmark year, conditioning on $\mathbf{X}$ and $\mathbf{Z}$.

[6] For these purposes, the unexpected departure (e.g., death) of a nonworking household member constitutes a per capita income shock.

are shocks to permanent income. For example, the receipt of a large, unanticipated inheritance prior to retirement would produce a significant contemporaneous consumption shock while at the same time elevating the level of wealth at retirement. This would tend to induce a positive correlation between the wealth ratio and the consumption growth rate prior to retirement (but not after retirement), biasing our procedure in favor of the theories that stress a connection between growth rates of consumption and wealth. Unless the estimated relation differs markedly before and after retirement, the magnitude of the bias is probably small.

In theory, one could remove bias created by the potential endogeneity of retirement wealth through the use of appropriate instrumental variables, for example, by predicting wealth at retirement based on information obtained prior to the household's earliest consumption observation. In practice, the available instruments, such as the value of the primary home and business income are, at best, only moderately good predictors of retirement wealth, so this procedure involves considerable loss of precision. Nevertheless, we use these instrumental variables to check the reliability of our main findings.

Finally, we consider the possibility that the timing of retirement is stochastic, and that premature retirement produces negative consumption innovations. As explained in Section I, subsection A.2, this would create a negative correlation between retirement wealth and the magnitude of the decline in consumption at retirement. To distinguish between the various hypotheses discussed in Section I, subsection A, it is important to determine whether those with low retirement wealth experience larger declines in consumption at retirement, even when the timing of retirement is properly anticipated. To do this, we must remove the effects of stochastic retirement.

Because data limitations force us to construct predictions of retirement based on relatively few instruments, we must simplify our model somewhat. In particular, assume that

$$\xi(t, \mathbf{X_i}) = \beta^* + (\mathbf{X_i}\boldsymbol{\beta}^0)I(t = 0). \tag{11}$$

Note that (11) is a special case of (8); here we allow for a baseline consumption trend as well as a discontinuity at retirement.[7] It follows that

$$\psi(t, \mathbf{X_i}) = t\beta^* + (\mathbf{X_i}\boldsymbol{\beta}^0)I(t \geq 0). \tag{12}$$

Substituting (12) into (9), switching time subscripts from age relative to retirement (t) to absolute age ($s = r_i + t$), and separating out the predictable and unpredictable components of retirement, we obtain

$$\begin{aligned} \ln C_{is} = {} & \hat{\mu}_i + s\beta^* \\ & + (\mathbf{X_i}\boldsymbol{\beta}^0)\Pr(s \geq r_i|\mathbf{Y_{is}}) \\ & + \boldsymbol{\Omega}\mathbf{Z_{is}} + \upsilon_{is}, \end{aligned} \tag{13}$$

where $\hat{\mu}_i \equiv \mu_i - r_i\beta^*$. We express the probability of being retired at age s, $\Pr(s \geq r_i|\mathbf{Y_{is}})$, as a function of deterministic household characteristics $\mathbf{Y_{is}}$. The error term is defined as follows:

$$\begin{aligned} \upsilon_{is} = {} & (\mathbf{X_i}\boldsymbol{\beta})[I(s \geq r_i) \\ & - \Pr(s \geq r_i|\mathbf{Y_{is}})] + \eta_{is}. \end{aligned} \tag{14}$$

Because the probability of retirement is not correlated with the transformed consumption error term, we can obtain consistent estimates by treating (13) as a standard regression equation. Identification of the baseline retirement effect depends on the fact that the retirement hazard function is nonlinear in age (spiking at ages 62 and 65), so we also add a quadratic age term.

Besides substituting the probability of retirement for the retirement dummy variable, as in equation (13), we also modify our specification by replacing wealth at retirement (in the vector $\mathbf{X_i}$) with a measure of wealth that is not contaminated by surprises about the timing of retirement. Intuitively, we would like to measure wealth at some standardized age that is sufficiently advanced for individuals to have accumulated a significant fraction of their retirement resources, but early enough to precede retire-

[7] As we will see, the data support the simplifying assumptions that the baseline consumption trend is independent of household characteristics, and that it is the same before and after retirement.

ment for that vast majority of households. Unfortunately, we do not have access to this information. Instead, we use the data on wealth that we do have, and adjust it to remove the effects of differences in age and retirement. In particular, we regress wealth on age, age-squared, a retirement dummy variable, and years retired (treating our sample as a cross section). Using the coefficients from this regression, we then adjust each household's wealth to a common age and work status. This is equivalent to using the residuals from the wealth regression as measures of "abnormal" wealth. By construction, this measure of wealth is orthogonal to information concerning retirement. However, it is still related to a wide range of idiosyncratic, unobserved characteristics that contribute to the unexplained variation in wealth at retirement.

III. Data

The primary data sample for our analysis consists of the set of all households surveyed in the Panel Study of Income Dynamics (PSID) with a transition to retirement between the years 1978 and 1990. We define nonretired households to be those with at least one member (head or spouse) working more than 1,500 hours annually. We define a household to be retired if no member works more than 500 hours annually in the current year, or in any subsequent year for which data are available. Naturally, some individuals made the transition from nonretired status to retired status over the course of several years, during which time one or more member worked part-time (between 1,500 and 500 hours).[8] We restricted the sample to households with transition periods of less than five years.[9] For the remaining observations, the variable t is set equal to -1 in the last year in which the household is nonretired, and equal to $+1$ in the first year in which the household is retired;[10] transition period data are excluded from our analysis. In the course of our empirical investigation, we controlled for lengthy transitions by including a dummy variable to identify households that spent more than two years (but fewer than five years) in transition to retirement; however, the coefficients of this variable were rarely significant and generally not large.

The total sample used in this analysis generally includes in excess of 3,500 observations on 430 households, with the specific number depending on the regression specification. The samples are generally unbalanced because later-retiring households are observed for fewer years after retirement, and more years before retirement, than early-retiring households. This does not, however, appear to drive our findings, as similar results are obtained when the regressions are estimated using subsamples consisting of balanced panels.

Unfortunately, the PSID does not contain ideal data on consumption. Although many past researchers have used food consumption to proxy for total consumption, Skinner's (1987) analysis of the Consumer Expenditure Survey (CEX) indicates that superior measures are available. Specifically, by using additional information on consumption reported in both the PSID and the CEX, such as the composition of food expenditures, utility payments, value of the house, and car ownership, one can increase the predictive power of the PSID consumption index threefold.[11] We use this approach below with a more restricted set of consumption indicators—food at home, food away from home (excluding meals at work or school), and the imputed or actual rental value of one's residence (utilities and autos were reported only sporadically)—although we also discuss regression results for each component separately. In certain years, the PSID did not collect information

[8] Households may have shifted from part time back to full time as well during this transition period. See Alan Gustman and Thomas Steinmeier, 1984.

[9] Thirty-nine percent of households made the transition to retirement in one year, 73 percent took no more than two years, 80 percent took no more than three years, and 85 percent took no more than four years. Thus, the restriction excludes roughly 15 percent of the potential sample. Of these, roughly half made the transition in five or six years.

[10] Note that for someone who made the transition from full-time to part-time work in 1982, and from part-time to fully retired in 1984, the year prior to retirement would be 1981 and the year after retirement would be 1984.

[11] In other words, using the Consumer Expenditure Survey data, the R^2 of a regression of total consumption on total food consumption was 0.26; with the additional components of consumption, the R^2 rose to 0.78. This is not surprising given the importance of rental or owner-occupied housing expenses in a typical budget.

on food consumption, and we were forced to exclude these observations from the sample.[12]

As mentioned previously, equations (7) and (9) allow the shape of the consumption profile to depend on the household's earnings replacement rate and retirement savings, which are summarized by the vector $\mathbf{X_i}$. To provide functional flexibility, we use dummy variables that indicate the household's position in the sample distribution of each variable.[13] Specifically, we divide our sample into four equally sized quartiles based on the ratio of nonasset income (total pension, social security, transfer, and earned income) for the first three years postretirement to nonasset income in the last three prior to retirement.[14] All measures of income are after-tax, where the household's tax rate is determined by taking the ratio of federal taxes paid by the head and spouse to the total income of the head and spouse. It is important to emphasize that we define earnings replacement in terms of ratios. We find, for example, that the fourth (or highest) quartile includes both high income households with generous postretirement compensation packages, as well as low income households with high social security replacement rates.

Similarly, we also divide our sample into four equally sized quartiles based on the ratio of wealth in the year prior to retirement to average preretirement (year $t = -3$ through $t = -1$) nonasset after-tax income. Recognizing that there is some disagreement in the literature on retirement saving concerning the appropriate measure of wealth [compare, e.g., Eric M. Engen et al. (1996) and James M. Poterba et al. (1996)], we estimated separate specifications for total wealth and financial wealth; generally, the results are quite similar.

Unfortunately, since the PSID collected comprehensive information on the components of wealth only in 1984 and 1989, we typically do not directly observe wealth in the year prior to retirement. In such cases, we extrapolate retirement wealth by applying the intertemporal budget constraint. Using observed wealth in either 1984 or 1989 (whichever is closer to retirement) along with estimates of consumption and measures of money income, we backcast (or project) wealth inductively according to the equation

$$W_{t-1} = \frac{W_t - Y_{t-1} + C_{t-1}}{1 + r}, \tag{15}$$

using a real interest rate (r) of 4 percent. Alternative approaches to imputing wealth in the year of retirement, such as using unadjusted wealth or adjusting observed wealth to the age of retirement using regression estimates on age and age-squared, yielded similar results.

Equation (7) also permits the consumption profile to depend on a vector of other time-varying household characteristics $\mathbf{Z_{it}}$. For the specifications reported here, this vector includes demographic variables that are likely to affect consumption, such as family size, disability status and gender of the household head, and marital status. The coefficient on female headship, for example, is thus identified by the difference between the effects on consumption of a husband's death and a wife's death. Twenty-seven percent of household heads report disabilities, 17 percent are female, and 68 percent are married.

Additional summary statistics for the sample appear in Table 1. There is substantial variation in income replacement rates, wealth ratios, and the level of wealth. These numbers are suggestive of the wide heterogeneity in retirement preparation among the retirement-aged population. The second-to-last row of Table 1 shows the change in (log) average consumption between the two years prior to retirement and the two years postretirement. The average change was -0.14, whereas the median decline was -0.12. These summary statistics mask substantial variation. The standard deviation of the change in log consumption was 0.42; the decline exceeded 30 percent for 23 percent of the

[12] We also dropped 28 observations for which the respondent reported zero food consumption, either at or away from home.

[13] For the results reported in the text, we use variables that measure the household's position in the sample distribution, rather than in the (weighted) population distribution. We obtain similar results when we define our variables in terms of population distributions. As we mention later, we also obtain similar results for specifications involving linear and piecewise linear functions of earnings replacement rates and retirement saving.

[14] Note that we exclude asset income from the numerator and denominator of the ratio.

TABLE 1—SUMMARY STATISTICS

Variable names	Mean	Median	Standard deviation	Minimum	Maximum
Age	62.1	62	5.4	45	80
Income replacement rate	0.63	0.60	0.32	0.02	2.87
Financial wealth–income ratio	2.33	0.88	5.80	−6.24	63.51
Total wealth–income ratio	3.91	2.36	6.45	−6.10	65.86
Log consumption (1984 dollars)	9.57	9.63	5.19	5.45	11.50
Difference in log consumption $(t_{-1/-2} - t_{+1/+2})$	−0.14	−0.12	0.42	−2.30	1.80
Family size (does not include spouse)	1.83	1	1.36	1	14

TABLE 2A—PROPORTIONAL DISTRIBUTION OF SAMPLE BY INCOME REPLACEMENT AND WEALTH QUARTILE

	Income Q1	Income Q2	Income Q3	Income Q4	Total
Wealth Q1	5.7	6.2	6.6	6.6	25.1
Wealth Q2	4.8	8.0	5.5	6.6	24.9
Wealth Q3	7.1	4.4	7.8	5.5	24.8
Wealth Q4	7.6	6.4	5.0	6.2	25.2
Total	25.2	25.0	24.9	24.9	100.0

TABLE 2B—AVERAGE AGE AT RETIREMENT BY INCOME REPLACEMENT AND WEALTH QUARTILE

	Income Q1	Income Q2	Income Q3	Income Q4	Total
Wealth Q1	59.5	60.4	60.0	61.1	60.3
Wealth Q2	59.1	61.3	61.3	60.8	60.8
Wealth Q3	60.0	61.7	61.8	60.4	61.0
Wealth Q4	59.7	61.9	61.6	61.7	61.1
Total	59.6	61.3	61.2	61.0	60.8

Note: Q1 denotes the lowest quartile and Q4 denotes the highest.

sample, whereas 12 percent of households experienced declines in excess of 40 percent.

Table 2A shows the joint distribution of the sample over the wealth ratio quartiles and income replacement quartiles. Under the life-cycle consumption-smoothing hypothesis, we would expect to observe a negative relation between retirement income replacement rates and retirement wealth. However, there is no evidence of this pattern. Only 24.1 percent of the sample falls into the Southwest/Northeast diagonal, compared with 27.7 percent for the Northwest/Southeast diagonal. Both figures are close to the sample frequency (25 percent) that one would observe if the observations were distributed randomly across the 16 cells. There may, of course, be a variety of subtle explanations for this pattern; taken by itself, it does not justify strong inferences concerning behavior.

Table 2B shows average retirement ages for each of the wealth ratio and income replacement quartiles. Differences in retirement age across these groups are small, and no systematic relations are apparent. This does not lend much credence to the view that the observed variation in wealth is in part attributable to unexpected developments affecting the timing of retirement. In particular, though early retirees are most likely to have entered retirement unexpectedly, they do not appear to have accumulated less for retirement, relative to their incomes.

IV. Basic Results

Table 3 presents estimates of our first-differenced specification [equation (7), with ξ defined as in equation (8)]. The first row describes the shape of the consumption profile for households in the lowest income replacement and wealth ratio quartiles (henceforth, the "benchmark"

TABLE 3—CONSUMPTION SHIFTS AND GROWTH RATES, PRE- AND POSTRETIREMENT: TOTAL CONSUMPTION

	Preretirement log consumption growth	First-year change in log consumption	Two-year change in log consumption	Postretirement log consumption growth
Benchmark (quartile 1 for income and wealth ratios)	−0.024 (0.012)	−0.240 (0.064)	−0.566 (0.078)	0.011 (0.029)
Wealth Ratio Quartile:				
Quartile 2	−0.029	−0.153	−0.291*	−0.018
Quartile 3	−0.015	−0.096*	−0.298*	−0.026
Quartile 4	−0.017	−0.094*	−0.211*	−0.035
p-value for equality of coefficients $F(3, 435)$	0.619	0.064	0.000	0.423
Income Replacement Quartile:				
Quartile 2	−0.027	−0.187	−0.338*	−0.006
Quartile 3	−0.025	−0.168	−0.355*	0.018
Quartile 4	−0.031	−0.052*	−0.199*	0.007
p-value for equality of coefficients $F(3, 435)$	0.924	0.017	0.000	0.849

Notes: $N = 3{,}262$. All significance tests use robust standard errors clustered by household (436). Additional variables are family size, marital status, disability, female widower, and a dummy variable for whether the household was working part-time for 3–4 years prior to full retirement.

* Denotes hypothesis that the quartile growth rate (or jump) differs from the benchmark is rejected at the 5-percent level in a two-tailed test.

group). There is a modest negative trend in consumption prior to retirement (−0.024). There is a substantial and significant drop within the first two years of retirement (−0.566), with slightly less than half of the decline occurring in the first year. Plainly, the impact of retirement on consumption is not instantaneous. The estimated postretirement consumption growth rate (0.01), though positive, is small and statistically insignificant.

The second section of Table 3 contains coefficients that describe the consumption profiles of individuals in the second, third, and fourth wealth ratio quartiles. There is virtually no variation in the growth rate of consumption across wealth quartiles either before retirement or after retirement; the significance value for the joint hypothesis of equality across all four wealth quartiles is 0.619 prior to retirement and 0.423 after retirement.[15] In contrast, the results in Table 3 indicate that there are large differences across wealth ratio quartiles in the size of the consumption discontinuity at retirement. Generally, a higher wealth ratio is associated with a smaller decline in consumption; in the top wealth quartile (and bottom income quartile), the estimated decline in consumption is −0.211, less than half the decline in the bottom wealth quartile (−0.566). The associated coefficients are individually significant at high levels of confidence, and one rejects the hypothesis of equal discontinuities across wealth ratio quartiles at the 99.9 percent level of confidence.

Figure 4 depicts consumption paths for households in each of the four total wealth ratio quartiles (assuming that the household falls into the first income replacement quartile), as well as for households in the top wealth ratio and income replacement quartiles. The horizontal axis

[15] As mentioned in Section III, shocks to permanent income just before retirement may generate a spurious relation between wealth at retirement and preretirement consumption growth. To examine this possibility, we implemented two alternative procedures: (i) we predicted wealth at retirement based on information obtained before each household's earliest consumption observation, and substituted predicted wealth for actual wealth; (ii) we confined attention to households for whom we observed wealth prior to retirement, used only subsequent consumption observations, and substituted observed wealth for retirement wealth. We found no evidence of a significant relation between wealth accumulation and preretirement consumption growth in either case.

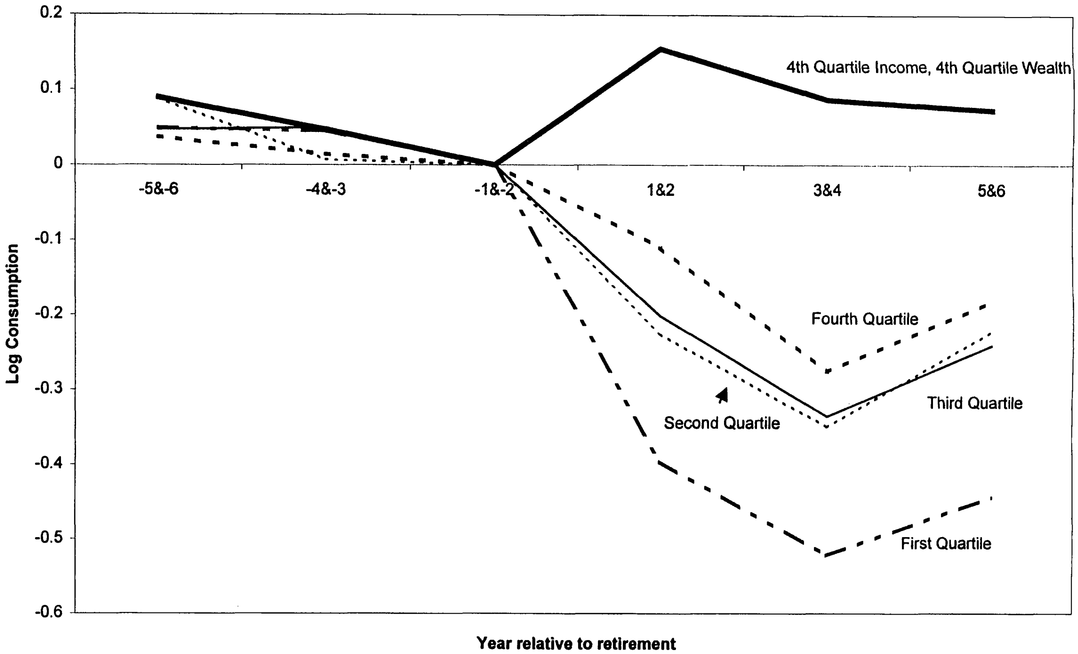

FIGURE 4. CHANGE IN CONSUMPTION AT RETIREMENT, BY WEALTH QUARTILE

measures years relative to retirement (e.g., $-5/-6$ refers to the fifth and sixth years before retirement). The vertical axis measures the log of normalized consumption. We define normalized consumption as the ratio of a household's consumption during any given time period to its consumption during a benchmark period (by arbitrary convention, the first and second years before retirement). This figure is based on the more flexible, less parsimonious, nonparametric fixed-effects specification discussed in Section II [equation (9), with ψ defined as in equation (10)]. Appendix Table A1 contains the associated parameter estimates. Once again, one sees that consumption growth rates do not differ systematically or significantly across wealth ratio quartiles either before or after retirement. Formally, one cannot reject the joint hypothesis that age-specific consumption growth rates are the same for all wealth quartiles both before and after retirement ($p = 0.82$).[16] However, the figure does exhibit large differences in consumption discontinuities at the time of retirement.

Differences in the shape of consumption profiles across income replacement quartiles are also intriguing. In the lower portion of Table 3, we see large and highly significant differences across these groups in the size of the consumption discontinuity at retirement. Higher income replacement rates are associated with smaller declines in consumption at retirement, and one can reject the hypothesis of an equal discontinuity across income replacement quartiles at the 99.9-percent level of confidence.[17] Figure 5 depicts consumption profiles for households in each of the four income replacement quartiles (assuming that the household falls into the first wealth ratio quartile), as well as for households in the top wealth ratio and income replacement quartiles. Like Figure 4, this figure is based on our more flexible, less parsimonious, nonparametric fixed-effects

[16] The effects of the discontinuities at retirement persist for many years; we obtain similar results when we use data including up to 10 postretirement years.

[17] Notably, these differences remain even when one restricts attention to households that actually engage in nontrivial saving (i.e., those in the top three wealth ratio quartiles).

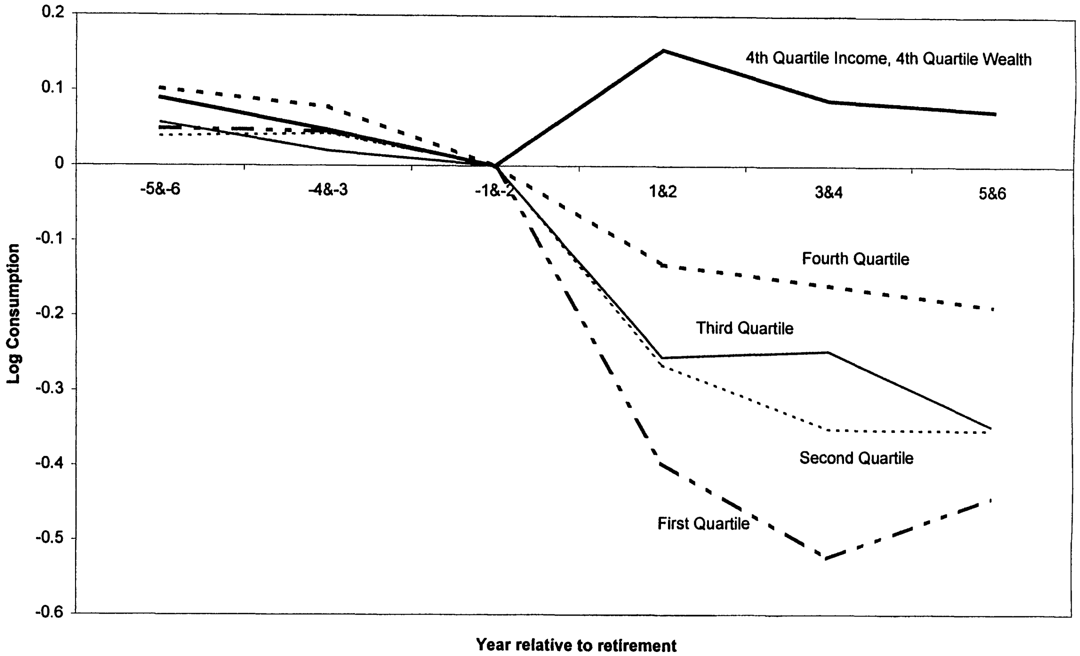

FIGURE 5. CHANGE IN CONSUMPTION AT RETIREMENT, BY INCOME QUARTILE

specification (Appendix Table A1). The figure clearly shows much steeper declines in consumption at retirement for households with lower income replacement rates. In combination with the data in Table 1, our nonparametric estimates imply that 30.4 percent of households reduce consumption by at least 35 percentage points within three to four years of retirement.

At first, the pattern noted in the preceding paragraph may appear to be reconcilable with models of consumption smoothing: because the regression also controls for the household's retirement wealth ratio, it effectively rules out smoothing through the endogenous adjustment of personal saving. However, recall from Table 2A that there is virtually no correlation between a household's positions in the wealth ratio and income replacement distributions. Consequently, even when wealth ratio variables are excluded from the regressions in Table 3, one still observes significantly larger discontinuities in consumption for households in lower income replacement quartiles. (We exclude the regression to conserve space.) As discussed in Section I, subsection B, this relation could also be attributable, at least in principle, to sample selection problems. However, as we have noted, it is possible to test for the most natural version of this problem by examining the composition of consumption, which we do in Section V.

Table 3 also indicates that the slope of the consumption profile does not differ much across income replacement quartiles either before or after retirement; the significance values of the F-tests for equality of the coefficients are 0.924 preretirement and 0.849 postretirement. The fact that those with the lowest income replacement rates exhibit virtually no attempt to anticipate the dramatic fall in consumption at retirement by scaling back on consumption prior to retirement (Table 3 and Figure 5) casts doubt on conventional models of consumption smoothing at retirement. Similarly, the absence of any correlation between consumption growth rates and income replacements rates is not generally consistent with models in which individuals who discount the future relatively little tend to save more, and to self-select into jobs with pension coverage (e.g., Gale, 1998).

The preceding results are based on flexible, nonparametric specifications of the relation between changes in consumption and measures of retirement wealth and postretirement income replacement rates. Simple parametric specifications yield similar results.[18]

V. The Composition of Consumption

As noted in Section I, there are a number of factors that could in principle account for the existence of a consumption discontinuity at retirement, as well as for correlations between the size of this discontinuity and variables such as wealth or income replacement rates. In most cases, theory also implies that these patterns should be confined to (or at least more pronounced for) particular kinds of expenditure categories (i.e., work-related expenses and substitutes for leisure). Thus, one can distinguish between theories at a more refined level by examining disaggregated measures of consumption.

Although the PSID is not ideally suited for this task, one can disaggregate somewhat by analyzing expenditure patterns separately for food consumed at home and away from home (where the latter category excludes meals consumed at work or school). If, as seems likely, home cooking is complementary with leisure and not work-related, then spending on food consumed at home should not decline after retirement. Moreover, variation in work-related expenses should produce a correlation between wealth and consumption only for potentially work-related spending categories, which presumably excludes food consumed at home. Provided that home cooking is at least less work-related and more highly complementary with leisure than restaurant meals, the decline in spending at retirement should be less pronounced for food consumed at home than for food consumed away from home.

Tables 4 and 5 contain estimates of our first-differenced specification [equation (7), with ξ defined as in equation (8)] for, respectively, food consumed away from home and food consumed at home.[19] In each case, the patterns are broadly similar to those noted in Table 3. However, for food consumed away from home (Table 4), the magnitude of the changes at retirement are considerably less pronounced. Indeed, one cannot reject the hypotheses that the discontinuity at retirement is the same across all wealth ratio quartiles ($p = 0.256$) and all income replacement quartiles ($p = 0.672$). Table 5 reveals that food consumed at home declines by a larger amount at retirement (-0.764 for households in the lowest wealth and income quartiles). The magnitude of this discontinuity declines monotonically with the household's wealth ratio quartile, and with its income replacement quartile. Moreover, we strongly reject the hypotheses that the discontinuity is the same across all wealth quartiles (99-percent level of confidence) and across all income replacement quartiles (99.9-percent level of confidence). These patterns are difficult to reconcile with the hypotheses that either the cessation of work-related expenses or substitution between market expenditures and leisure explain the sharp decline in spending at retirement.[20]

Notably, in Table 5, we also reject the hypothesis that the postretirement rate of growth for food consumed at home is the same across wealth ratio quartiles. Specifically, the pattern of coefficients suggests that consumption tends to decline at a more rapid rate after retirement for those who reached retirement with greater wealth. The same pattern was present in Table

[18] For example, we estimated specifications in which we replaced the wealth ratio and income replacement quartile dummies with the household's percentile rank for each of these variables. Interactions with preretirement and postretirement dummies were jointly insignificant ($p = 0.452$). With respect to the first year retirement effect, the coefficient of the wealth ratio variable was 0.00213 (t-statistic of 2.83), and the coefficient of the income replacement ratio variable was 0.00225 (t-statistic of 2.99). With respect to the second year retirement effect, the coefficient of the wealth ratio variable was 0.00220 (t-statistic of 2.68), and the coefficient of the income replacement ratio variable was 0.00191 (t-statistic of 2.41). We tested the implicit linearity assumption by estimating another specification with splines at the median values of both variables. Based on the appropriate F test, we failed to reject the linear specification ($p = 0.338$).

[19] We obtain similar results for the fixed-effects specification.

[20] There are of course some limitations of this test. At retirement, households might switch from high-cost prepared food purchased at the supermarket to less expensive basic foods prepared at home. However, the degree of substitution to low-cost items would have to be quite extreme (and among just the low wealth and income quartiles) to generate the pattern seen in the data.

TABLE 4—CONSUMPTION SHIFTS AND GROWTH RATES, PRE- AND POSTRETIREMENT: FOOD AWAY FROM HOME

	Preretirement log consumption growth	First-year change in log consumption	Two-year change in log consumption	Postretirement log consumption growth
Benchmark (quartile 1 for income and wealth ratios)	0.003	−0.140	−0.502	0.122
	(0.049)	(0.198)	(0.265)	(0.101)
Wealth Ratio Quartile:				
Quartile 2	0.021	−0.100	−0.300	0.018
Quartile 3	−0.042	0.170	−0.085	0.080
Quartile 4	0.014	−0.064	−0.249	0.100
p-value for equality of coefficients $F(3, 381)$	0.297	0.144	0.256	0.564
Income Replacement Quartile:				
Quartile 2	−0.025	−0.153	−0.334	0.015
Quartile 3	0.036	−0.171	−0.336	0.048
Quartile 4	0.038	−0.035	−0.261	0.072
p-value for equality of coefficients $F(3, 381)$	0.282	0.788	0.672	0.595

Notes: $N = 2{,}396$. All significance tests use robust standard errors clustered by household (382). Additional variables are family size, marital status, disability, female widower, and a dummy variable for whether the household was working part-time for 3–4 years prior to full retirement.

* Denotes hypothesis that the quartile growth rate (or jump) differs from the benchmark is rejected at the 5-percent level in a two-tailed test.

TABLE 5—CONSUMPTION SHIFTS AND GROWTH RATES, PRE- AND POSTRETIREMENT: FOOD AT HOME

	Preretirement log consumption growth	First-year change in log consumption	Two-year change in log consumption	Postretirement log consumption growth
Benchmark (quartile 1 for income and wealth ratios)	−0.043	−0.349	−0.764	0.022
	(0.018)	(0.098)	(0.137)	(0.028)
Wealth Ratio Quartile:				
Quartile 2	−0.054	−0.177	−0.514*	−0.079*
Quartile 3	−0.035	−0.156	−0.466*	−0.050
Quartile 4	−0.062	−0.138	−0.380*	−0.083*
p-value for equality of coefficients $F(3, 435)$	0.196	0.123	0.007	0.018
Income Replacement Quartile:				
Quartile 2	−0.040	−0.355	−0.482*	−0.022
Quartile 3	−0.055	−0.227	−0.363*	0.037
Quartile 4	−0.057	−0.112	−0.240*	0.014
p-value for equality of coefficients $F(3, 435)$	0.552	0.006	0.000	0.477

Notes: $N = 3{,}248$. All significance tests use robust standard errors clustered by household (436). Additional variables are family size, marital status, disability, female widower, and a dummy variable for whether the household was working part-time for 3–4 years prior to full retirement.

* Denotes hypothesis that the quartile growth rate (or jump) differs from the benchmark is rejected at the 5-percent level in a two-tailed test.

TABLE 6—CONDITIONAL BUDGET SHARE REGRESSIONS, CONSUMER EXPENDITURE SURVEY, 1982–1989

Dependent variable	Adult clothing	Transportation	Fuel	Food away from home	Food at home
Average Budget Share (Percent)	*2.50*	*14.85*	*4.39*	*4.22*	*14.63*
Retired (yes = 1) ×	−0.448	−1.665	−0.704	−0.982	1.893
Wealth quartile 1	(6.1)	(4.6)	(6.4)	(7.0)	(9.4)
Retired (yes = 1) ×	−0.539	−1.634	−0.930	−1.306	1.367
Wealth quartile 2	(6.5)	(4.0)	(7.5)	(8.2)	(6.0)
Retired (yes = 1) ×	−0.227	−1.506	−0.487	−0.534	0.040
Wealth quartile 3	(1.9)	(2.6)	(2.8)	(2.4)	(0.1)
Retired (yes = 1) ×	−0.243	−2.827	−0.665	−0.229	−0.301
Wealth quartile 4	(3.0)	(7.1)	(5.5)	(1.5)	(1.4)
Married	0.418	−2.228	0.464	−3.105	3.797
	(5.5)	(6.1)	(4.1)	(21.3)	(18.2)
Single female head	1.110	−2.800	−1.369	−3.268	−0.734
	(14.5)	(7.4)	(12.0)	(22.4)	(3.5)
Any children ≤ 18	−0.204	1.268	0.389	−0.627	1.685
yrs? (yes = 1)	(1.5)	(1.9)	(2.0)	(2.5)	(4.6)
Number of children	−0.229	−0.833	0.084	−0.331	1.783
	(3.2)	(2.4)	(0.8)	(2.4)	(9.1)
Log(total expenditures)	0.843	8.595	−0.475	1.227	−8.094
	(19.0)	(39.1)	(7.1)	(14.5)	(66.4)
R^2	0.12	0.27	0.15	0.12	0.37

Notes: Five-year dummy variables for age and individual year dummy variables included in all regressions. The sample size is 10,260. All dependent variables are percentages of total expenditures. Absolute values of *t*-statistics are in parentheses.

3, but the differences across wealth ratio quartiles were statistically insignificant. Recall that the life cycle considerations discussed in Section I, subsection A.1 predict precisely the opposite pattern: those with higher wealth should have higher consumption growth rates.

Because the PSID collected data on relatively few expenditure categories, it is important to determine whether our results are more broadly representative. We therefore extend our investigation by analyzing Julie A. Nelson's (1994) extract of the 1980–1989 Consumer Expenditure Survey (CEX), which contains merged and annualized quarterly consumption data. We restrict our attention to 1982 through 1989 because of concerns about poor data quality in the start-up years. Observations are dropped if the household's data are missing for any quarter, or if the appropriate fields indicate that income data are of poor quality.

The CEX data are, for practical purposes, a series of cross sections. The survey does not provide a measure of each household's wealth at retirement, or of changes in income after retirement. We therefore use the conditional commodity demand approach (see, e.g., Martin Browning and Costas Meghir, 1991) to study the effect of retirement on relative budget shares. In effect, our object is to determine whether labor market status (here, retirement) has independent predictive power for the budget shares of particular expenditure categories (work-related expenses and leisure complements), controlling for prices and total expenditures. In that we are not interested in estimating the complete budget system, we adopt a simplified version of the Browning–Meghir specification, wherein we simply include dummy variables to allow for variation in prices across years. We would expect that individuals with higher-than-average tastes for work might also have higher-than-average tastes for goods complementary with work; thus we would expect our ordinary least-squares (OLS) estimates to be upper bounds on the "true" impact of retirement on commodities such as transportation expenses, adult clothing, and fuel.

Table 6 contains average budget shares and parameter estimates for a simple conditional budget share model. The model explains relative expenditures on several categories of goods that are plausibly complements to working (adult clothing, transportation, fuel, and food away from home including meals at work), as well as one non-work-related expenditure category that is presumably a complement to leisure (food consumed at home). The model includes dummy variables for age and year (not reported) as well as the demographic variables listed in the table.[21]

A negative partial correlation between retirement and potentially work-related expenses would not by itself indicate that such factors can explain the *variation* in wealth across the population. To conclude that lower wealth households save less because they anticipate larger declines in work-related expenses, one would have to find that households in lower wealth quartiles experience *larger* declines in potentially work-related spending at retirement. To investigate this possibility, we interact the retirement dummy variable with wealth-to-income quartile dummies constructed from the limited wealth data (on stocks, bonds, saving bonds, and checking accounts) contained in the Consumer Expenditure Survey.[22] More than one-third of the sample has a missing value for at least one wealth category; all missing wealth values were set to zero. Thus we recognize that these wealth quartiles are measured with error.

As is clear from the coefficients in the second through fifth rows of Table 6, retirement is indeed associated with lower budget shares for goods that are potentially complementary to work.[23] However, focusing on *differences* in the impact of retirement on budget shares across wealth groups, there is relatively little variation by wealth in the share of the budget spent on adult clothing (a fall of 0.25 percentage points for the high wealth group compared to a fall of 0.45 percentage points for the low wealth group) or on fuel (a decline of 0.67 percentage points for the high wealth group and a decline of 0.70 percentage points for the low wealth group). Indeed, the transportation share declines by *more* among the top wealth group than among the bottom wealth group. These findings are inconsistent with the hypothesis that wealth varies significantly across households as a result of the anticipation of differential declines in work-related clothing, fuel, or transportation expenses.

The final column of Table 6 suggests that the budget share devoted to food consumed at home does increase after retirement, and that this increase is larger for households with less wealth. This may at first appear consistent with the hypothesis that households shift to expenditures on home-production activities postretirement. Recall, however, that the equation explains *budget shares*. Because retirement is, in this sample, associated with a decline of 0.23 in the log of total consumption, the equation predicts a substantial decline in spending on food consumed at home after retirement.[24] Our results are therefore inconsistent with the prediction that the decline in consumption at retirement should be confined to work-related expenses and leisure substitutes.

VI. Removing the "News" Associated with Retirement

For the reasons discussed in Section I, the preceding empirical patterns are difficult to reconcile with many explanations for the observed variation in wealth based on standard models of life-cycle optimization. The evidence presented so far does not, however, exclude the possibility that the variation in retirement wealth and the associated variation in the size of the consumption discontinuity at retirement are both attributable, at least in part, to unexpected events that affect the timing of retirement.

To examine this possibility, we investigate the manner in which consumption responds to

[21] We also estimated a model closer in spirit to the Almost Ideal Demand System (see, e.g., Deaton and John Muellbauer, 1980 pp. 75–84); results were generally similar with interaction terms between all of the explanatory variables and the log of total consumption expenditures, but much less stable.

[22] The wealth quartile dummy variables indicate the household's position within the wealth-to-income distribution after adjusting for differences in retirement status, age, and marital status.

[23] This result is consistent with similar tests in Banks et al. (1998).

[24] The budget share for food consumed at home would increase by 1.86 percentage points at retirement for the lowest wealth group if total expenditures remained constant. Given that food at home is income inelastic (the coefficient of the log of total expenditures is -8.09), the budget share rises slightly as total expenditures decline. However, these effects are much smaller, individually and collectively, than the overall decline in consumption.

TABLE 7—SECOND-STAGE CONSUMPTION REGRESSIONS

Variable name	A Total wealth	B Financial wealth	C Total wealth, demographics excluded	D Total wealth, retirement age > 60
Probability of retirement	−0.685	−0.557	−0.627	−0.606
	(4.3)	(3.7)	(3.6)	(3.4)
Wealth quartile 2 × Pr(Ret)	0.287	0.280	0.261	0.235
	(2.0)	(1.9)	(1.7)	(1.5)
Wealth quartile 3 × Pr(Ret)	0.416	0.291	0.451	0.426
	(2.9)	(2.1)	(2.9)	(2.7)
Wealth quartile 4 × Pr(Ret)	0.594	0.430	0.596	0.597
	(4.0)	(3.0)	(3.8)	(3.6)
Income quartile 2 × Pr(Ret)	0.261	0.208	0.294	0.269
	(1.8)	(1.4)	(1.9)	(1.7)
Income quartile 3 × Pr(Ret)	0.298	0.266	0.341	0.286
	(2.2)	(2.0)	(2.4)	(1.9)
Income quartile 4 × Pr(Ret)	0.344	0.312	0.465	0.284
	(2.7)	(2.4)	(3.4)	(2.0)
Family size	0.079	0.080		0.084
	(5.3)	(5.2)		(6.8)
Marital status	0.147	0.148		0.136
	(6.3)	(6.3)		(5.1)
Female head	−0.231	−0.227		−0.263
	(2.8)	(2.7)		(2.4)
Age	−0.092	−0.093	−0.094	−0.094
	(3.7)	(3.6)	(3.8)	(2.6)
$Age^2/1{,}000$	0.587	0.584	0.497	0.587
	(2.8)	(2.8)	(2.4)	(2.1)
Constant	8.200	8.218	8.837	8.331
	(10.2)	(10.1)	(11.6)	(6.8)
R^2	0.76	0.76	0.74	0.76
(Observations)	(4,817)	(4,817)	(4,817)	(3,609)

Notes: Dependent variable is log of consumption. Household-specific fixed effects included in second-stage regression. First-stage probits are run separately for each age group; all z statistics (in parentheses) are based on bootstrapped standard errors.

predictable events that affect the probability of retirement. In particular, to identify the effect on consumption of predictable retirement, we exploit the fact that the retirement hazard function varies sharply with age, spiking at ages 62 and 65 (see Banks et al., 1998, for a similar identification strategy).

As explained in Section II, we remove the effects of unexpected retirement through a two-step procedure. First, we estimate simple probit specifications explaining retirement status as a function of education, family size, gender of household head, and marital status as independent variables. We fit separate models for each integer age from 54 through 70, and in each instance we augment the data set to include all observations on all households, irrespective of whether they retired between 1978 and 1990. We then introduce the predicted probability of retirement into a second-stage consumption regression, as in equation (13). For the second stage, we calculate wealth ratio quartiles using a measure of "abnormal" wealth instead of wealth at retirement. We bootstrap the two-step procedure (using 1,000 replications) to obtain standard errors.

Table 7 presents estimates of the second-stage regression. The results in column A, like those in Table 3, control for family size, marital status, the gender of the household head, and age.[25] The estimated impact of predicted retirement for those in the lowest wealth and income quartiles is −0.685, with a z-statistic of 4.3. The magnitude of this effect declines monotonically

[25] For the regressions in Table 7, we omit disability on the grounds that it may serve as a proxy for retirement. Note also that we allow the baseline consumption trend to depend on both age and age-squared.

across wealth ratio quartiles: it is −0.398 in the second quartile, −0.269 in the third quartile, and −0.091 in the fourth quartile.[26] Similarly, the magnitude of the effect declines monotonically across income replacement quartiles: it is −0.424 in the second quartile, −0.387 in the third quartile, and −0.341 in the fourth quartile. Regardless of whether one examines wealth or income, the differences between the effects for the lowest quartile and each of the higher quartiles are statistically significant at conventional levels of confidence. Thus, even when we remove the effects of unexpected retirement, the size of the consumption discontinuity is still strongly related to wealth and income—if anything, the estimated patterns are even more striking. These estimates, combined with information from Table 1, imply that 31 percent of households reduce their consumption by at least 35 percentage points at retirement.[27]

In columns B through D of Table 7, we examine the robustness of our findings with respect to several changes in specification. For column B, we use financial wealth in place of total wealth. For column C, we exclude controls for demographic characteristics on the grounds that changes in these variables may be associated with shocks to needs and/or permanent income. For column D, we confine attention to individuals retiring after age 60 on the grounds that shocks to permanent income are probably largest for early retirees.[28] In each case, we find that our central results are highly robust.

[26] One obtains these numbers by adding the coefficient for the appropriate interaction term to the effect for the lowest wealth and income quartile (e.g., −0.398 = −0.685 + 0.287).

[27] Note that these figures are essentially identical to the ones reported in Section IV for our nonparametric OLS estimates. The implied fraction of households suffering a substantial decline in consumption is much larger than the figure attributed to us by Engen et al. (1999 p. 135). They used coefficient estimates from a two-stage specification in an earlier version of this paper (NBER Working Paper No. 6227, October 1997) that was not designed to distinguish between age effects and the impact of a baseline retirement trend. For this reason, their interpretation of our results is problematic.

[28] When an individual is forced to retire unexpectedly before age 60, the present discounted value of lost earnings may be quite large. The financial impact is even greater for those who retire early for health reasons (because of limitations on Medicare eligibility) and for those whose pension plans heavily penalize early retirement.

VII. Conclusion

In this paper, we have found that consumption growth rates near retirement do not vary systematically with retirement wealth. Thus there is no indication that heterogeneity in pure rates of time preference, subjective survival probabilities, income uncertainty, or tastes for precaution—all of which should manifest themselves through systematic differences in consumption growth rates—play a role in determining the distribution of retirement savings. We have found a pronounced discontinuity in consumption at retirement, with the size of the discontinuity negatively correlated with retirement savings and income replacement rates. However, none of these phenomena is confined to work-related expenses or leisure substitutes. The empirical evidence therefore casts doubt on theories that rely on differences in relative tastes for leisure, home production, or work-related expenses to explain the variation in wealth at retirement.

Likewise, differences in retirement wealth for households with similar income and pension profiles (including retirement ages) could in principle result from heterogeneity in planned retirement ages, provided that realized retirement ages are affected to some degree by unanticipated events. Although this factor is also consistent with the negative correlation between the consumption discontinuity and retirement savings, it cannot account for the fact that this same pattern remains readily apparent even when we remove the effects of unpredictable events that affect the timing of retirement. Moreover, whereas variation in the strength of bequest motives may contribute to differences in wealth accumulation, it fails to explain the strong negative correlation between retirement savings and the magnitude of the consumption discontinuity at retirement.

We are also unable to attribute differences in retirement accumulation to variation in the shape of the income/pension profile. Households with lower income replacement rates have larger consumption discontinuities at retirement. Contrary to the central tenets of life-cycle theory, there is little evidence that households use savings to smooth the effects on consumption of predictable income discontinuities.

These findings are difficult to interpret in the context of the life-cycle model. Although it may be possible to formulate some model of rational life-cycle planning that would account for our findings, in our view, the empirical patterns in this paper are more easily explained if one steps outside the framework of rational, farsighted optimization. If, for example, households follow heuristic rules of thumb to determine saving prior to retirement, and if they take stock of their financial situation and make adjustments at retirement (so that the adequacy of saving is "news"), then one would expect to observe the patterns documented in this paper. A similar conclusion may follow from theories of "mental accounting," in which individuals differ in the extent to which they can exercise self-discipline over the urge to spend current income, and/or from models with dynamically inconsistent decision makers.

Appendix

Table A1—Regression for Total Consumption

		Interaction terms with year relative to retirement				
		−5 and −6	−4 and −3	+1 and +2	+3 and +4	+5 and +6
Benchmark (1st quartile, wealth and income)		0.049	0.045	−0.396	−0.522	−0.444
		(1.0)	(1.2)	(6.9)	(7.7)	(4.3)
Increment—2nd wealth quartile		0.039	−0.038	0.170	0.172	0.221
		(0.9)	(1.1)	(3.4)	(2.9)	(2.7)
Increment—3rd wealth quartile		−0.002	0.004	0.194	0.186	0.202
		(0.1)	(0.1)	(3.5)	(2.9)	(2.3)
Increment—4th wealth quartile		−0.012	−0.031	0.286	0.246	0.261
		(0.3)	(0.8)	(5.1)	(4.0)	(2.9)
Increment—2nd income quartile		−0.010	−0.003	0.131	0.171	0.091
		(0.3)	(0.1)	(2.4)	(2.6)	(0.9)
Increment—3rd income quartile		0.008	−0.026	0.141	0.275	0.096
		(0.2)	(0.7)	(2.5)	(4.2)	(1.0)
Increment—4th income quartile		0.053	0.032	0.265	0.363	0.256
		(1.2)	(0.8)	(4.8)	(5.8)	(2.8)
Part-time work		−0.026	−0.061	−0.047	−0.051	−0.063
		(0.5)	(1.4)	(0.9)	(0.7)	(0.9)
Family size	0.044					
	(3.3)					
Female head (widowhood)	−0.137					
	(2.1)					
Married	0.077					
	(3.2)					
Disabled	−0.047					
	(2.6)					

Notes: Sample size = 3,980. Estimates correct for household-specific effects with robust standard errors (clustered by household). $R^2 = 0.75$. Dependent variable is log total consumption. Absolute values of t-statistics are in parentheses. The top row shows benchmark coefficients for bottom wealth and income quartiles. The "increment" for each wealth ratio and income replacement quartile measures log consumption *relative* to the benchmark group.

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