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The nature of precautionary wealth

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Abstract

This paper uses the Panel Study of Income Dynamics to provide some of the first direct evidence that wealth is systematically higher for consumers with predictably greater income uncertainty. However, the apparent pattern of precautionary wealth is not consistent with a standard parameterization of the life cycle model in which consumers are patient enough to begin saving for retirement early in life; wealth is estimated to be far less sensitive to uncertainty than implied by that model. Instead, our results suggest that over most of their working life time, consumers behave in accordance with the 'buffer-stock' models of saving described in Carroll (1992, 1997) or Deaton (1991), in which consumers hold wealth principally to insulate consumption against near-term fluctuations in income.

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1. Introduction

Recent advances in the theory of precautionary saving[1] have sparked interest in whether precautionary saving is an empirically important phenomenon. This paper uses the Panel Study of Income Dynamics (PSID) to provide direct evidence that wealth is higher for households with greater income uncertainty. We also show, however, that the pattern of precautionary wealth holding does not appear to be consistent with a standard life cycle optimization problem in which retirement saving is important at the early stages of working life. Our empirical results are consistent with a version of the life cycle model like that in Carroll (1992, 1997), in which consumers engage in 'buffer-stock' saving behavior during most of their working lifetimes and only begin to save for retirement around the age of 50.

Our empirical approach is to construct estimates of labor income uncertainty using the PSID and then to investigate the relationship between uncertainty and wealth. We first decompose uncertainty into a variance of shocks to permanent (life time) income and a variance of shocks to transitory income. Controlling for demographic effects and for the level of income, our empirical results indicate that net worth depends importantly on the degree of both transitory and permanent income uncertainty.

We then use the empirical results to discriminate between competing versions of the life cycle/permanent income hypothesis (LC/PIH) model of consumption under uncertainty. We use numerical techniques to solve the problem under the parametric assumptions used by Hubbard, Skinner and Zeldes (1994, henceforth HSZ), which imply that consumers begin saving for retirement early in life. We show that under the HSZ parametric assumptions, wealth holdings should be highly sensitive to the degree of uncertainty in permanent income. The model predicts a regression coefficient on the variance of permanent shocks at least ten times larger than the coefficients estimated in our empirical work.

We next demonstrate that when the model is solved under alternative parameter values which imply that consumers engage in 'buffer-stock' saving over most of the life time (Carroll (1992, 1997), the sensitivity of wealth to uncertainty is far lower – enough lower, in fact, to be consistent with our empirical estimates. This is because buffer-stock savers have an effective 'horizon' of only a few years (using 'horizon' in the loose sense of Friedman (1957)), while consumers actively engaged in retirement saving have an effective horizon of the remainder of their life time. We also demonstrate the intuitive result that in a buffer-stock model of saving, wealth becomes less sensitive to uncertainty as consumers become more impatient. We find that under the Carroll (1992, 1997) choices for parameters

[1] Most notably by Zeldes (1989), Kimball (1990a, 1990b, 1991), Caballero (1990, 1991) and Carroll (1992, 1997).

other than the rate of time preference, the buffer-stock model can exactly match our empirical results if the rate of time preference is approximately 11% annually.

The paper is organized as follows. In Section 2, we discuss the various methods that others have used to study precautionary saving and then explain the relative advantages of our approach. In Section 3, we develop our technique for estimating the magnitude of permanent and transitory shocks to income and present these estimates by occupation, education, and industry group. The econometric results relating wealth to income uncertainty are given in Section 4. In Section 5, we present the numerical solutions to the life cycle model with income uncertainty under the HSZ and the Carroll (1992, 1997) parameterizations, and we develop our methodology for estimating a time preference rate which would be consistent with our empirical results. Section 6 provides the concluding remarks.

2. Empirical evidence on the existence of precautionary saving

It would be a considerable understatement to say that no consensus has emerged from the empirical literature on precautionary saving. In a widely cited early article, Skinner (1988) found that saving was actually *lower* than average for two groups of households (farmers and the self-employed) who face higher-than-normal income uncertainty. Guiso et al. (1992) found that consumption was only slightly lower and wealth slightly higher for Italian consumers with a greater subjective variance of next year's income. Dynan (1993) estimated the coefficient of relative prudence and found it to be small and insignificantly different from zero. (In the absence of prudence, there is no precautionary motive for saving.) But other studies have found evidence of a very strong precautionary saving motive. Dardanoni (1991) examined consumption and income data for British households and found that average consumption across occupation and industry groups was negatively related to the within-group variance of income; he estimated that around 60% of saving is due to precautionary motives. Carroll (1994) found that income uncertainty was statistically and quantitatively important in regressions of current consumption on current income, future income, and uncertainty. Finally, Carroll and Samwick (1995b) estimated that about a third of household wealth is attributable to the fact that some households face greater uncertainty than others.

An advantage of our approach over most of these previous studies is that all but Guiso et al. (1992) and Carroll and Samwick (1995b) relate uncertainty to current consumption rather than to the stock of wealth. As a theoretical proposition, the appropriate response to greater uncertainty is to hold more wealth; it is not necessarily to depress consumption forever. In the steady state of a buffer-stock saving model, for instance, average consumption will approximately equal average income for both low and high uncertainty consumers, thus maintaining the average buffer stock constant at the optimal level for each group. Thus, across consumers already holding the optimal buffer stock, there will be *no* apparent relation between current consumption or the current saving rate and the uncertainty of income. However, consumers with greater uncertainty will desire a larger buffer stock, so there *will* be a relationship between the level of wealth and the degree of uncertainty.[2]

A further criticism of the previous studies is that no one used an entirely appropriate measure of income uncertainty. Using the panel dimension of income observations in the PSID, we are able to make direct estimates of the variance of innovations to permanent income for each household – the theoretically correct measure of uncertainty for the models used by Dardanoni (1991) and Guiso et al. (1992). Furthermore, we can make a distinction between innovations to permanent income and transitory shocks to income that the more strictly utility-based measures used by Carroll (1994) and Carroll and Samwick (1995b) do not distinguish.[3]

3. Estimating the variances of transitory and permanent shocks

Our data set from the PSID contains income data for the years 1981–1987 for approximately 4,000 households, along with data on household wealth in 1984. In this section, we define our measure of income and develop a method for decomposing innovations into transitory and permanent components. We also present empirical estimates of income uncertainty by occupation, education, industry group, and age in the PSID and describe the consequences of measurement error and other data problems for our estimates of uncertainty. We use only data from 1981–87 period, on the view that income uncertainty measures constructed using data from these years will correspond reasonably well to the true uncertainty for the mid-point year of 1984, when wealth is measured.[4]

Our income measure (total household non-capital income) includes (a) labor income of the head, spouse, and other household members; (b) disability payments, welfare payments, and other forms of transfer income (including food stamps and in-kind transfers); (c) unemployment insurance and social

[2] Until the optimal buffer stock is achieved, there will be a relation between consumption and uncertainty – the consumer facing higher uncertainty will initially have to depress consumption more in order to build up the larger stock of wealth.

[3] For a more comprehensive discussion of the literature, see Carroll and Samwick (1995a).

[4] We also performed the analysis using data from 1976–1987, with results similar to those we report.

security payments; and (d) almost all other kinds of income except direct capital income.[5]

For expositional clarity, we will begin by discussing a model of the income process that is slightly simpler than the one we actually estimate using the PSID data.[6] The logarithm of permanent income p_t is assumed to follow a random walk with drift

$$p_t = g_t + p_{t-1} + \eta_t, \tag{1}$$

where g_t represents predictable growth due to life cycle aging and to overall aggregate productivity growth and η_t is the shock to permanent income in period t. The log of current income is given by the log of permanent income plus a transitory error term:

$$y_t = p_t + \varepsilon_t. \tag{2}$$

Assume that the errors ε and η are white noise and uncorrelated with each other at all leads and lags. Remove the predictable component of income growth g_t and rewrite (1) as

$$p_t = p_{t-1} + \eta_t. \tag{3}$$

Define a d-year income difference as

$$\begin{aligned} r_d &= y_{t+d} - y_t \\ &= p_{t+d} + \varepsilon_{t+d} - p_t - \varepsilon_t, \end{aligned} \tag{4}$$

where the second equality is obtained by substituting (2) into the first equality. Substituting (3) into (4) recursively yields

$$r_d = \{\eta_{t+1} + \eta_{t+2} + \ \dots \ + \eta_{t+d}\} + \varepsilon_{t+d} - \varepsilon_t. \tag{5}$$

Finally, obtain the d-year variance as the second moment of the right hand side of Eq. (5):

$$\mathrm{Var}(r_d) = d\sigma_\eta^2 + 2\sigma_\varepsilon^2, \tag{6}$$

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[5] Ideally, income would be defined net of taxes as well as transfer payments because a progressive income tax structure reduces the variation in disposable income available to finance consumption [see Kimball and Mankiw (1989)]. However, federal and state tax payments are not directly measured in the data set and any tax calculations based only on the reported income data would certainly introduce considerable further measurement error. Furthermore, Pechman (1985) asserts that, when all forms of taxation are considered, and the effects of tax shelters (widespread in this period) and other tax avoidance strategies are figured in, the tax burden at this time was roughly flat. A flat income tax would affect the level of log income, but would not affect the variance.

[6] Our methodology here is a generalization of that of Hall and Mishkin (1982).

where σ_η^2 and σ_ε^2 are the variances of the permanent and transitory shocks to income, respectively. For households whose income behaved according to this model, we could estimate $\mathrm{Var}(r_d)$ for each household i by $v_{id} = r_{id}^2$, where r_{id} is constructed as in (5). This would yield an unbiased estimate of $\mathrm{Var}(r_{id})$ if the mean of r_{id} is zero, which corresponds to an assumption that there are no individual-specific growth rates for income (other than those predictable by occupation, education, industry, and other personal characteristics).[7] For each household, we could use any two v_d's of different lengths to solve for σ_η^2 and σ_ε^2. For example, for any $d > 1$, one way of estimating σ_η^2 for a given household is simply to take (we suppress the subscript i for clearer notation):

$$s_\eta^2 = v_d - v_{d-1}, \tag{7}$$

$$s_e^2 = \frac{v_{d-1} - (d-1)s_\eta^2}{2} \tag{8}$$

and it is easy to show that $\mathrm{E}(s_\eta^2) = \sigma_\eta^2$ and $\mathrm{E}(s_\varepsilon^2) = \sigma_\varepsilon^2$. The appendix shows how to derive a more efficient estimate of σ_ε^2 and σ_η^2 by optimally combining all available v_d's which can be constructed over the entire sample period. The appendix also shows that it is simple to generalize the method to allow for serial correlation in the transitory component of order MA(q).

We allow for serial correlation in the transitory term of order MA(2), because the literature on the time series process for household labor income [see MaCurdy (1982), Abowd and Card (1989)] has found no evidence for a transitory component of order greater than MA(2). To remove the predictable component of income growth, we divide actual income by the predicted value from a regression of log income on age, occupation, education, industry, household demographic variables, and age-interaction terms. The predicted values are also adjusted for economy-wide growth in income so that the average normalized income variable shows no trend over the sample period.

Table 1 presents the resulting estimates of income uncertainty for the occupation, education, and industry groups we distinguish in the PSID. Breakdowns by age are also shown. The measure in the first column is an overall measure of uncertainty over the 1981–87 sample period, i.e., the square of the difference in (detrended) income between 1981 and 1987, divided by six to yield an annual rate. In the notation introduced above, for group j (e.g. for occupation group 1), $u_j = \text{mean}\,(r_{i6}^2/6)$ where the mean is taken across all households in that group. The comparisons across groups are plausible: farmers, service workers, and self-employed managers have high uncertainty; laborers, clerical workers, and managers have average uncertainty; professionals and craftsmen have the least

[7] MaCurdy (1982) tested this proposition for labor income of the household head in the PSID and found no evidence against it.

Table 1
Income uncertainty by occupation, education, industry and age

Group	(1) Variance of the annual innovation in income	(2) Estimated variance of the permanent component	(3) Estimated variance of the transitory component	Percent of sample ($N = 1325$)
Full sample	0.0355	0.0217	0.0440	100.00
	(0.0018)	(0.0029)	(0.0055)	
Occupation				
Professional and technical workers	0.0292	0.0172	0.0331	22.3
	(0.0039)	(0.0062)	(0.0116)	
Managers (not self-employed)	0.0321	0.0180	0.0357	13.2
	(0.0050)	(0.0081)	(0.0151)	
Self-employed managers	0.0453	0.0165	0.0926	4.2
	(0.0089)	(0.0143)	(0.0267)	
Clerical and sales workers	0.0388	0.0235	0.0361	14.3
	(0.0048)	(0.0078)	(0.0145)	
Craftsmen	0.0285	0.0175	0.0432	20.0
	(0.0041)	(0.0066)	(0.0123)	
Operatives and laborers	0.0403	0.0299	0.0458	17.1
	(0.0044)	(0.0071)	(0.0133)	
Farmers and farm laborers	0.0660	0.0450	0.1016	2.3
	(0.0119)	(0.0192)	(0.0359)	
Service workers	0.0490	0.0189	0.0611	6.5
	(0.0071)	(0.0116)	(0.0215)	
p-value	0.008	0.287	0.885	
Education				
0–8 Grades	0.0382	0.0190	0.0894	4.6
	(0.0085)	(0.0137)	(0.0256)	
9–12 Grades	0.0459	0.0214	0.0658	10.6
	(0.0056)	(0.0090)	(0.0168)	
High school diploma	0.0383	0.0277	0.0431	18.0
	(0.0043)	(0.0069)	(0.0129)	
Some college, no degree	0.0353	0.0238	0.0342	38.8
	(0.0029)	(0.0047)	(0.0088)	
College degree	0.0284	0.0146	0.0385	18.9
	(0.0042)	(0.0068)	(0.0126)	
Some advanced education	0.0324	0.0115	0.0500	9.1
	(0.0061)	(0.0097)	(0.0181)	
p-value	0.218	0.680	0.272	

Table 1 (continued)

Group	(1) Variance of the annual innovation in income	(2) Estimated variance of the permanent component	(3) Estimated variance of the transitory component	Percent of sample ($N = 1325$)
Industry				
Agriculture, forestry, fishing	0.0596	0.0401	0.0794	3.6
	(0.0096)	(0.0155)	(0.0288)	
Mining	0.0431	0.0389	− 0.0018	1.1
	(0.0172)	(0.0277)	(0.0516)	
Construction	0.0452	0.0313	0.0494	6.5
	(0.0072)	(0.0116)	(0.0215)	
Manufacturing	0.0317	0.0249	0.0275	28.9
	(0.0034)	(0.0055)	(0.0102)	
Transportation, communication, and utilities	0.0302	0.0111	0.0603	10.8
	(0.0056)	(0.0090)	(0.0167)	
Wholesale and retail trade	0.0408	0.0231	0.0489	14.3
	(0.0048)	(0.0078)	(0.0145)	
Finance, insurance and real estate	0.0286	0.0118	0.0510	4.9
	(0.0082)	(0.0133)	(0.0248)	
Business and repair services	0.0256	0.0037	0.0662	3.7
	(0.0095)	(0.0153)	(0.0285)	
Personal services	0.0655	0.0200	0.1040	1.7
	(0.0147)	(0.0229)	(0.0426)	
Entertainment and recreation services	0.0329	0.0107	0.0892	0.5
	(0.0251)	(0.0405)	(0.0755)	
Professional and related services	0.0392	0.0216	0.0469	16.4
	(0.0045)	(0.0072)	(0.0136)	
Public administration	0.0218	0.0119	0.0218	7.6
	(0.0066)	(0.0107)	(0.0200)	
p-value	0.018	0.990	0.500	
Age				
25–30	0.0305	0.0205	0.0418	23.9
	(0.0037)	(0.0060)	(0.0121)	
31–35	0.0305	0.0170	0.0411	24.7
	(0.0037)	(0.0059)	(0.0104)	
36–40	0.0331	0.0219	0.0422	15.1

Table 1 (continued)

Group	(1) Variance of the annual innovation in income	(2) Estimated variance of the permanent component	(3) Estimated variance of the transitory component	Percent of sample ($N = 1325$)
41–45	0.0326	0.0030	0.0780	10.6
	(0.0056)	(0.0090)	(0.0168)	
46–50	0.0433	0.0248	0.0404	11.3
	(0.0054)	(0.0087)	(0.0163)	
51–57	0.0511	0.0389	0.0320	14.4
	(0.0048)	(0.0073)	(0.0144)	
p-value	0.005	0.073	0.426	

Notes:

1) Group designations pertain to the head of household in the beginning year of the sample (1981).
2) Column (1) is the variance of the unexplained component of the difference in income between the beginning and end of the sample period (1981–1987), divided by the length of the sample period.
3) Columns (2) and (3) are the average values of the estimated variance of permanent or or transitory income for all households in the specified group.
4) Standard errors are given in parentheses.

uncertainty. For the households with between nine years of education and a college degree, uncertainty declines with education. Among industry groups that comprise 10% or more of the sample, workers in the manufacturing and utilities sectors face less uncertainty than those in trade or professional services. Additionally, workers in public administration face comparatively little income uncertainty by this measure. The final breakdown is by age group; we found little overall variation in uncertainty by age for consumers in the age group of 25–45, but somewhat higher uncertainty for older consumers.

For each of the occupation/industry/education/age breakdowns we present p-values for the hypothesis that uncertainty is identical across groups. For the overall measure of uncertainty, the p-values indicate highly significant differences in the u_j across occupation (p-value 0.008), industry (p-value 0.018), and age (p-value 0.005) groups, but insignificant differences by education (p-value 0.218).

The decomposition of income uncertainty into the variances of permanent and transitory shocks in the second and third columns is also plausible, although the high p-values indicate that these cross-group differences are usually not statistically significant at conventional levels. The potential importance of decomposing income variance into permanent and transitory components is clear, for example, in a comparison of laborers to service workers. While the latter have a higher overall measure of uncertainty, this is due to a very large variance of transitory shocks; the magnitude of permanent shocks is substantially higher for the laborers.

On its face, the finding that the permanent variances are not significantly different by occupation, education, industry, and age group appears to bode ill for using these variables as instruments for permanent variance. However, the p-value tests performed in Table 1 are not proper formal tests for instrument validity, because the full instrument set allows for differences in the age profile of uncertainty by group, and accounts for the effects of covariation among groups. We present the results from the appropriate formal tests of instrument validity, which our full instrument set passes easily, as shown below.

Measurement error in income often causes important problems in attempts to measure uncertainty. However, since we intend to focus primarily on the coefficient on permanent uncertainty, measurement error may not cause major problems for our analysis. To see why, suppose measurement error is of the commonly specified type:

$$y_t^{\mathrm{r}} = y_t^{\mathrm{true}} + u_t,$$

where y_t^{r} signifies reported income and y_t^{true} signifies true income. Then if u_t is white noise and the variance of the error term, $\mathrm{Var}(u_t) = \sigma_u^2$, is uncorrelated with personal characteristics, measurement error will have no effect on our estimate of permanent uncertainty. This can be noticed from

$$\begin{aligned} E(s_\eta^2) &= \mathrm{Var}(r_d) - \mathrm{Var}(r_{d-1}) \\ &= d\cdot\sigma_\eta^2 + 2\cdot(\sigma_\varepsilon^2 + \sigma_u^2) - [(d-1)\cdot\sigma_\eta^2 + 2\cdot(\sigma_\varepsilon^2 + \sigma_u^2)]\,. \\ &= \sigma_\eta^2\,. \end{aligned} \tag{9}$$

Measurement error in current income is therefore treated by our procedure in the same way as transitory income shocks. Thus, the proof in the Appendix that our estimate of the permanent variance is robust to MA(2) serial correlation in the transitory shock also demonstrates that, in order for our estimate of the permanent variance to be corrupted by measurement error, the measurement error term would have to be more persistent than MA(2). Naturally, measurement error would increase our estimate of the level of transitory variance:

$$E(s_e^2) = \sigma_e^2 + \sigma_u^2 \tag{10}$$

but if σ_u^2 is the same for everyone, the only effect of measurement error would be to raise the estimate of everyone's transitory variance by the same amount, which would have no effect on our coefficient estimates.[8]

[8] One unfortunate property of our estimates of permanent and transitory variance is that they are negatively correlated across households. For a detailed discussion of the reasons, and for a proof that our instrumenting technique should eliminate this problem, see the working paper version of this paper, Carroll and Samwick (1995a).

4. How does wealth depend on uncertainty?

Three measures of wealth are considered in our empirical work. (i) Very liquid assets (VLA) are those assets which could be liquidated immediately. This includes bank account balances, money market funds, certificates of deposit, government savings bonds, mutual funds, and publicly traded stocks. (ii) Non-housing, non-business wealth (NHNBW) adds to VLA the net value of all other assets and liabilities not related to the primary residence or personally owned businesses. Such assets include non-government bonds, the cash value of whole life insurance, cars and other vehicles, secondary real estate, and other investments. In addition to loans or mortgages on any of these assets, NHNBW also deducts the balances on credit cards, student loans, outstanding medical and legal bills, and loans from relatives. Most of the components of this measure could be liquidated (or at least have the amount of equity altered) within a matter of weeks or months. (iii) Total net worth (NW) adds to NHNBW what are generally the most illiquid assets owned by households – equity in the primary residence and the net value of personally owned businesses.

With measures of wealth and uncertainty in hand, the next step is to regress household wealth on uncertainty in our sample of households. However, at the level of the individual household our direct measures of the variance of transitory and permanent shocks to income will be subject to enormous measurement error. Eqs. (9) and (10) show that *in expectation* our measures of uncertainty correspond to the true measures of uncertainty, but for any *individual* household our measures are a very noisy measure of the actual uncertainty faced, because our method essentially estimates two variance parameters from only seven income observations on the household's level of income. For a given household, our estimates of the variances can be represented as the true variances plus some household-specific error term. If we were to estimate an OLS regression of wealth on our constructed measures of household uncertainty, we would have a standard errors-in-variables problem: the estimated coefficients on the uncertainty terms would be biased toward zero with the magnitude of the bias proportional to the variance of the household-level measurement error. Since at the household level the variance of the error must be very large, the coefficients from an OLS regression would be severely biased.[9] We adopt the standard solution to such errors-in-variables problems, which is to estimate the equation using instrumental variables.

Because the life cycle model implies that wealth is also influenced by factors other than uncertainty, the regressions include demographic variables and

[9] In the OLS regressions we estimated, the coefficients on the uncertainty terms were in fact much smaller than in the instrumental variables specification.

a measure of household permanent income. The equation we estimate is

$$\log W = \alpha_0 + \alpha_1 s_\eta^2 + \alpha_2 s_\varepsilon^2 + \alpha_3 \log P + X\beta + u, \tag{11}$$

where P is an estimate of permanent income (defined here to be the average household income over the sample period) and X contains age and other demographic variables to control for predictable life cycle effects on wealth.[10]

Now we have to decide which variables are to be used as instruments for the income and uncertainty terms and which are to be included among our control variables X; to achieve identification, the instrument set must include variables that are not also in X. We chose to define X to include age, marital status, race, sex, and the number of children in the household, and to exclude occupation, education, and industry from X in order to identify the model.[11] These variables will be valid instruments only if, first, they have predictive power for income and uncertainty and, second, they have no predictive power for wealth *beyond* their influence through their ability to predict income and uncertainty. Bound et al. (1993) suggest the use of the partial R^2 and F-statistic on the excluded instruments in the first stage regression as rough guides to the quality of the IV estimates. The partial R^2s of the excluded instruments in the first stage regressions are 0.106 for the variance of permanent shocks and 0.078 for the variance of transitory shocks. The p-values for the F-test of joint significance of the instruments are 0.0001 and 0.0379, respectively.[12] Thus, our estimates of Eq. (11) should not suffer from the econometric problems first highlighted by Nelson and Startz (1990a, b) that can occur when the first-stage regressions perform poorly.[13]

The second stage regression results for Eq. (11) are presented in Table 2 for the sample of households 50 years old or younger for each of the three measures

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[10] Note that the coefficients in this equation all represent the long-run response of wealth to the independent variables. If, say, uncertainty were to increase suddenly for an individual consumer, wealth would not instantly jump to the value predicted by this equation; instead, it would move gradually in that direction. We should also mention here that we experimented fairly extensively with nonlinear terms in the measures of uncertainty and income, but did not find any nonlinear specifications that notably improved the fit over the linear specification.

[11] The traditional life cycle model suggests several other variables that might be related to saving. Among these are the expected date of death of the members of the househod; the expected income growth rate of the household; and the expected pension replacement rate for wages on retirement. We calculated measures of all three of these variables and experimented with including them among the controls, but none of these variables were systematically significant and none had a substantial impact on the estimated coefficients on the uncertainty variables of interest. For simplicity, therefore, we left them out of the final specification.

[12] These statistics correspond to the sample comprised of observations that report valid data for NHNBW. The VLA sample yielded partial R^2 of 0.0847 and 0.0732 and p-values of 0.0062 and 0.0488 for permanent and transitory variances, respectively. For the NW sample, the statistics were 0.1066, 0.0691, 0.0007 and 0.1086.

[13] See Staiger and Stock (1994) for an analysis of the effect of weak instruments on the second stage regression results. The test of instrument relevance based on the canonical correlations of the instrument set proposed by Hall et al. (1994) also rejected the null hypothesis for each sample with p-values below 0.001.

Table 2
Instrumental variables regressions of wealth on uncertainty for households aged 50 and under

	Very liquid assets	Non-housing, non-business wealth	Total net worth
Constant	− 26.33	− 14.86	− 12.00
	(3.49)	(2.70)	(2.51)
Permanent variance	12.09***	9.43***	8.64***
	(3.93)	(3.04)	(3.05)
Transitory variance	7.11***	3.88**	3.86**
	(2.20)	(1.90)	(1.78)
Permanent income	3.07	1.94	1.65
	(0.33)	(0.25)	(0.23)
Age	0.14	0.20	0.21
	(0.10)	(0.08)	(0.08)
$Age^2 \times 10^{-2}$	− 0.16	− 0.24	− 0.23
	(0.14)	(0.10)	(0.10)
Married	− 0.50	0.00	0.47
	(0.25)	(0.20)	(0.20)
Race	0.30	0.29	0.44
	(0.24)	(0.19)	(0.20)
Female	0.01	0.01	− 0.01
	(0.17)	(0.18)	(0.15)
Kids	− 0.24	− 0.11	− 0.04
	(0.06)	(0.05)	(0.05)
Overidentification test (p-value)	0.26	0.50	0.075
Number of observations	881	847	861

Notes:
1) Heteroskedasticity-robust standard errors are given in parentheses.
2) Instrumental variables used for the two income variances and permanent income are occupation, education, and industry dummies listed in Table 1, the occupation and education dummies interacted with Age and Age^2, and the demographic 'control' variables.
3) Demographic variables: Married = 1 if married, 0 otherwise; Race = 1 if white, 0 if non-white; Female = 1 if female head of household, 0 otherwise; Kids = number ofchildren under 18 in the household.
4) ***, **, and * denote statistical significance of uncertainty measures at the 1%, 5%, and 10% levels, respectively.

of wealth.[14] The key result is that for all three measures of wealth, the variances of both permanent and transitory shocks have positive effects on wealth. These estimates are significantly different from zero at the 1% (permanent) and 5% (transitory) levels for all the three measures of wealth. The fact that the coefficient

[14] We restrict the sample to those younger than 50 because the theoretical results presented below, and in Carroll (1997), suggest that the behavior of households younger than about 50 may be qualitatively different from the behavior of those older than 50. For details, see the theoretical analysis below.

estimates decline as the measure of wealth becomes more comprehensive means that the proportional effect of uncertainty on wealth declines as the measure of wealth becomes more comprehensive and less liquid. The fact that the proportional effect of uncertainty on wealth declines with the measure of wealth does not imply that greater uncertainty merely causes consumers to reshuffle wealth from less liquid to more liquid forms. In fact, the absolute effect of uncertainty on wealth increases as the measure of wealth becomes more comprehensive. For example, for a consumer at the median level of wealth, an increase in the standard deviation of permanent shocks from 5% to 10% would boost VLA by $2,139, but would boost NW by $13,627. So net worth increases by more than the increase in VLA despite the smaller coefficient in Table 2.

Our instrumental variables estimation procedure implicitly assumes that our instruments (occupation, education, industry, and their age interactions) are correlated with wealth only in-so-far as they are correlated with the variables for which we are instrumenting, uncertainty and permanent income. We test the validity of this assumption by considering the instrumental variables regressions in a generalized method of moments framework and using the heteroskedasticity-robust test of the overidentifying restrictions given in Hansen (1982). The p-values are reported at the bottom of the table. For VLA and NHNBW, the p-values are not small enough to reject the specification, but the p-value of 0.0750 for NW suggests that the occupation, education, or industry variables may have a direct effect on total net worth.[15,16,17]

[15] Despite these favorable OID tests, it of course remains possible that our instruments are related to wealth through channels other than their correlation with uncertainty and income. We chose occupation, education, and industry and their age-interactions as our instruments primarily because these are the variables used in previous work on the relationship of consumption to income and income growth [Carroll (1994), Carroll and Weil (1994)]. A thorough search of the data set for other variables that, *a priori*, we might have thought would be related to uncertainty found no variables that were able to explain much of the variation in uncertainty beyond that explained by occupation, education, and industry.

[16] If occupation group is added to the X variables, the overidentification restriction for NW is no longer rejected; the coefficients on the income uncertainty terms are reduced but are not significantly different (at the 5% level) from those reported in Table 2. However, the coefficients on the uncertainty terms for NW are no longer statistically significantly different from zero.

[17] A remaining caveat with the estimates in Table 2 is that, because wealth is measured in logs, households for whom reported net worth is negative (approximately 10% of the sample) have been omitted. In order to determine whether the bias from this censoring is empirically important, we estimated two-stage Heckit models in which the inverse Mills ratio from a probit of whether or not the observation reported strictly positive wealth was included as a regressor in the second stage regression. For each measure of wealth, the estimated coefficients on the uncertainty variables are not statistically different from the estimates in Table 2 and are still significantly different from zero at the 1% (permanent) and 5% (transitory) levels. For example, the coefficients in the NHNBW regression are 8.78 and 3.81 for permanent and transitory variances, respectively, compared to 9.43 and 3.88 in Table 2.

Table 3 presents a small sample of the many robustness checks we performed. The results from each check are summarized by the coefficients and t-statistics on the permanent and transitory variances (henceforth PVAR and TVAR). The first robustness test is for the effect of including the entire age range of consumers who were younger than 62 (the social security early retirement age) in our sample. The coefficient on PVAR generally increases and its statistical significance rises. The next test shows the results when we perform a median regression rather than a least squares regression. The coefficient on PVAR generally falls modestly, as does statistical significance, but there are no dramatic changes. The third set of tests involves relaxing our identifying assumption that education, industry, and occupation affect wealth only through their effect on uncertainty. We add the dummy variables for first education, then industry, and then occupation to the set of 'control' variables for wealth designated by X in Eq. (11), while keeping the dummy variables not added, and the age interactions, in the instrument set. Allowing education or industry to have a direct effect on wealth does not greatly change the results. However, when occupation dummies are added to the controls, the coefficients on PVAR and TVAR fall substantially and are no longer significant. Further investigation revealed that the two occupation groups mainly responsible for this effect are farmers and the self-employed. Both of these groups have relatively high uncertainty and relatively high wealth, and as a result they have a substantial impact on the coefficient estimates. The final set of robustness checks, therefore, estimates the equation when farmers and the self-employed are excluded from the sample. The coefficient estimates on PVAR for the VLA and NHNBW measures of wealth are about a third lower than when the farmers and self-employed are included in the sample, and both are statistically significant at only around the 10% level (as compared with better than 1% in the whole sample). The coefficient falls by more than half when total net worth is the dependent variable.

Our preferred interpretation of these findings is that occupation groupings provide some of the most important identifying variation in our sample, and removing that identifying variation reduces the information content of the data. It is worth noting here that the coefficients on PVAR and TVAR for the sample excluding the farmers and self-employed are not statistically significantly different from the coefficients when farmers and the self-employed are included.

Note that this evidence is precisely the opposite of what would be expected if the arguments of Friedman (1957) and Skinner (1988) about occupation and sample selection were correct. They argued that only the less risk-averse households would enter risky occupations, because less risk-averse households have a smaller precautionary saving motive; if this were true the efffect would be to bias coefficients on PVAR and TVAR down when occupation is used as an instrument. Instead, instrumenting occupation causes the estimated coefficients to rise and gain significance.

Table 3
Robustness checks

	Very liquid assets	Non-housing, non-business wealth	Total net worth
Baseline specification			
Permanent variance	12.09***	9.43***	8.64***
	(3.93)	(3.04)	(3.05)
Transitory variance	7.11***	3.88**	3.86**
	(2.20)	(1.90)	(1.78)
Include ages > 50			
Permanent variance	13.27***	10.28***	10.41***
	(4.22)	(3.16)	(3.35)
Transitory variance	6.60***	4.66**	5.28**
	(2.26)	(1.92)	(1.93)
Median regression			
Permanent variance	7.75**	9.09***	7.18***
	(3.82)	(3.48)	(2.06)
Transitory variance	3.94*	4.56**	2.07
	(2.16)	(2.19)	(1.35)
Education as control			
Permanent variance	11.08***	7.44**	8.97***
	(3.99)	(3.28)	(3.08)
Transitory variance	8.97***	3.88**	5.72***
	(2.66)	(1.90)	(2.04)
Industry as control			
Permanent variance	9.94**	5.28	5.68*
	(5.07)	(3.37)	(3.13)
Transitory variance	5.75**	0.95	2.10
	(2.95)	(2.28)	(2.08)
Occupation as control			
Permanent variance	4.78	7.19*	3.77
	(5.50)	(4.27)	(5.07)
Transitory variance	1.64	1.72	− 0.12
	(2.75)	(2.20)	(0.05)
Drop farmers, self-employed			
Permanent variance	8.90	6.99	3.39
	(5.55)	(4.13)	(4.14)
Transitory variance	3.81	2.03	0.47
	(2.75)	(1.92)	(0.24)

Notes:
1) The first three rows reproduce the relevant baseline results from Table 2 for easier comparison with the other results. (See notes to Table 2).
2) See text for detailed description of the experiments reported in this table.

Of course, it remains possible that these kinds of self-selection problems do exist in our sample so that the true coefficient is even larger than we estimate. There are other possible reasons our coefficients may be biased downward. One important reason is that our measures of uncertainty are based on pre-tax income. This is becaues the PSID does not collect income tax data from participants.[18] To the extent that a progressive tax code provides implicit insurance, our measures of uncertainty may exaggerate the true extent of uncertainty. In practice, the bias from this source is likely to be modest; Pechman (1985) estimated that when all the various tax loopholes and shelters that existed in this period were taken into account, and state and local taxes factored in, the tax burden was roughly flat in this time period. Further, even the nominal structure of the tax code was insufficiently progressive to bias our coefficients more than a maximum of around 30 or 40%.

5. Simulated results from a model of precautionary saving

We now wish to determine whether these empirical results are consistent with existing theories of precautionary saving. This section answers that question by solving a standard life cycle model with a precautionary saving motive and determining what coefficient values such a model would predict. We find that when the model is solved under parameter values like those used by Hubbard et al. (1994) it implies much greater response of wealth to uncertainty than we found, but when it is solved under parameter values which imply that consumers engage in 'buffer-stock' saving behavior until around age 50 it implies coefficient estimates roughly consistent with our empirical estimates. Finally, we show that our empirical results provide an implicit method for estimating consumers' time preference rates, because the model implies a monotonic relationship between the regression coefficient on uncertainty and the time preference rate. We then calculate the time preference rates implied by our empirical findings for a variety of configurations of parameter values.

5.1. The saving model

Our model of household saving is

$$\max_{C_t} \sum_{t=s}^{T} \beta^{t-s} u(C_t),$$

$$Y_t = P_t V_t,$$

[18] The PSID does contain a formula-based estimate of the taxes each household pays, but this is a very imperfect measure of actual taxes. We preferred not to add yet more measurement error to our estimate of income.

s.t. (12)

$$P_t = G_t P_{t-1} N_t,$$

$$X_t = R[X_{t-1} - C_{t-1}] + Y_t,$$

where gross wealth (what Deaton (1991) calls cash-on-hand) is $X_t = W_t + Y_t$, the sum of net wealth and current labor income (i.e. non-capital income, as discussed above); current labor income Y_t is given by permanent labor income P_t multiplied by a random mean-one transitory error V_t; permanent labor income P_t is equal to last period's permanent labor income multiplied by a random mean-one error N_t; the time preference factor is $\beta = 1/(1 + \delta)$ where δ is the time preference rate; and the utility function is of the CRRA form $u(c) = c^{(1-\rho)}/(1-\rho)$, where ρ is the coefficient of relative risk aversion. The structure of the income uncertainty follows Zeldes (1989) and Carroll (1992, 1997). The multiplicative transitory shock to income, V_t, is zero with probability π, corresponding to periods of unemployment when income is zero. If V_t is not zero, it is distributed lognormally. We assume that N_t, the multiplicative shock to permanent income, is also distributed lognormally, so that the log of permanent income follows a random walk with drift. The model is solved using standard numerical recursive dynamic stochastic programming methods. For details about the solution methods, see Carroll (1992, 1996, 1997).

This model differs from that of HSZ in several ways. First, HSZ directly impose liquidity constraints. In our model, consumers voluntarily choose never to borrow, because, with a positive probability of zero income events in each period, borrowing any finite sum induces a positive probability that consumption will be driven to zero in some future period (if enough bad income draws in a row arrive), and with CRRA utility, zero consumption results in infinitely negative utility. For readers uncomfortable with this logic, it will be easier to justify the zero-probability events as an alternative way to effectively put liquidity constraints in the model without introducing the additional computational difficulties that result from an explicit liquidity constraint. For further discussion of these issues, see Zeldes (1989) or Carroll (1992, 1997).

A second difference is that HSZ characterize the stochastic process for income as an AR(1) with a coefficient of 0.95 on lagged income, whereas our specification assumes that the log of permanent income has a unit root. We prefer our characterization, partly because we strongly believe that both transitory and permanent shocks exist, and partly because the model is much more difficult to solve with an AR(1) specification. HSZ assume that the standard deviation of the innovations to their income process is around 0.095; this is the analogue in their estimation to the variance of $\log(N)$ in our model. For simplicity, we will identify $\sigma_\eta = 0.1$ as the HSZ parameter value. We also set $\sigma_\varepsilon = 0.1$ for the transitory shock. These choices are also identical to the assumptions in Carroll (1992, 1997).

HSZ model two kinds of uncertainty neglected in our model – length-of-life risk and medical risk. However, they find that these two kinds of risk have fairly minor effects on saving behavior; the lion's share of precautionary saving in their model is induced by labor income risk. A final difference between this model and that of HSZ is that the HSZ model contains an elaborate and carefully calibrated description of the social insurance system.

Despite this catalog of differences between our model and that of HSZ, when our model is calibrated with parameter values similar to those used by HSZ, it produces similar results for mean and median wealth holding behavior. This is not surprising, because none of the differences between the two models should have a major impact on mean or median wealth: HSZ show that mortality risk and medical risk are unimportant; both models effectively prevent consumers from borrowing, although they do so in different ways; and the consumers at the bottom of the income distribution who are most affected by the social insurance system would not be holding much wealth anyway.

5.2. Results using the HSZ calibration

Turning now to the details, HSZ assumed an interest rate equal to the time preference rate at 3% annually; a coefficient of relative risk aversion $\rho = 3$; and a median age-income profile that can be approximated by one in which (real) income grows at around 2% annually until age 55, then declines at about 1% annually until retirement at age 65. At retirement, income drops to about 70% of its immediate pre-retirement level and remains constant thereafter.

The results from solving and simulating the model under the HSZ parameter values are depicted as the solid line in Fig. 1, which shows mean values for the ratio of net wealth to permanent labor income (henceforth, the net wealth ratio) by age. Consumers begin saving for retirement early in life and build up a large stock of assets over the course of a lifetime of saving.[19] These assets are then depleted after retirement, and consumers', assets dwindle to zero towards the end of their life.

The dashed line shows how the results change when the assumption about the standard deviation of the annual shock to permanent income is reduced from the HSZ baseline value of 0.1 to a value of 0.05. The effect on wealth accumulation is dramatic – wealth is much lower over the entire working lifetime. The

[19] The sharp increase in the graph at age 65 is due to the revaluation upon retirement of permanent income to 70% of its pre-retirement level, thereby increasing the ratio of net wealth to permanent income for the given stock of net wealth accumulated during the working years.

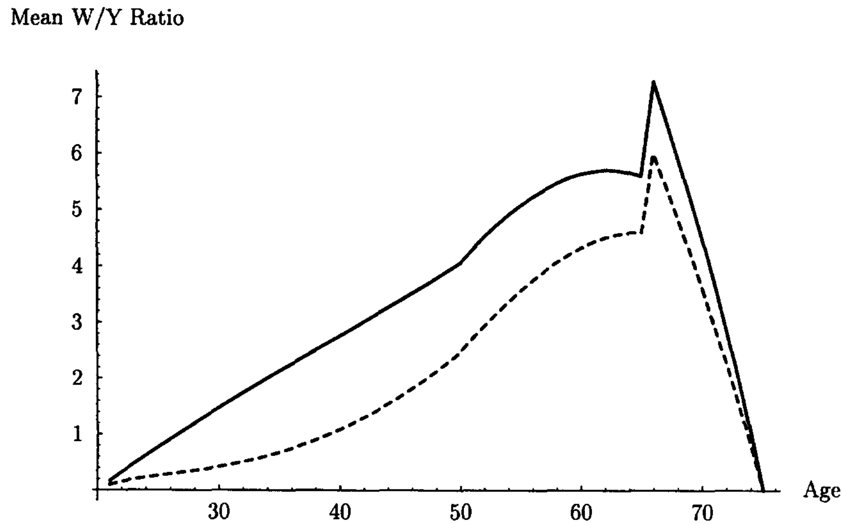

Fig. 1. Age–Wealth profile under HSZ parameters.

experiment just performed provides a way to approximate the coefficient that the model would predict in regressions like those estimated above. Consider a consumer at age 30. Under the baseline HSZ parameter values with $\sigma_\eta = 0.1$, the predicted mean net wealth ratio is 1.473. Under the alternative with $\sigma_\eta = 0.05$ and with all other parameter values the same, the predicted wealth ratio at 30 is 0.424. Hence, the coefficient on a regression of $\log(W)$ on σ_η^2 would yield a coefficient that can be approximated by $[\log(1.473)-\log(0.424)]/(0.1^2 - 0.05^2) \approx 166$.

The coefficients predicted by the model using the HSZ parameter values can be estimated in the same way for each age of life,[20] the results are shown in Fig. 2.[21] One inconvenient feature of the results is that the predicted coefficient on σ_η^2 is quite different at different ages. Due to the limited size of our empirical sample we were not able to estimate age-specific coefficients on σ_η^2. However, it

[20] The results are basically same when the experiment is changed to an increase of 0.05 in the standard deviation of shocks rather than a decrease, or when a smaller change in the standard deviation of uncertainty is considered.

[21] The upturn in the last couple of years of life represents consumption of the precautionary assets that are no longer needed because there is no future to save for. If we were to complicate the model slightly by having a steadily increasing mortality probability until certain death at, say, age 120, this upturn would be eliminated. Experimentation with such models suggested that they provide little additional insight for the question at hand, so we preferred to leave the model as simple as possible by keeping the date of death certain.

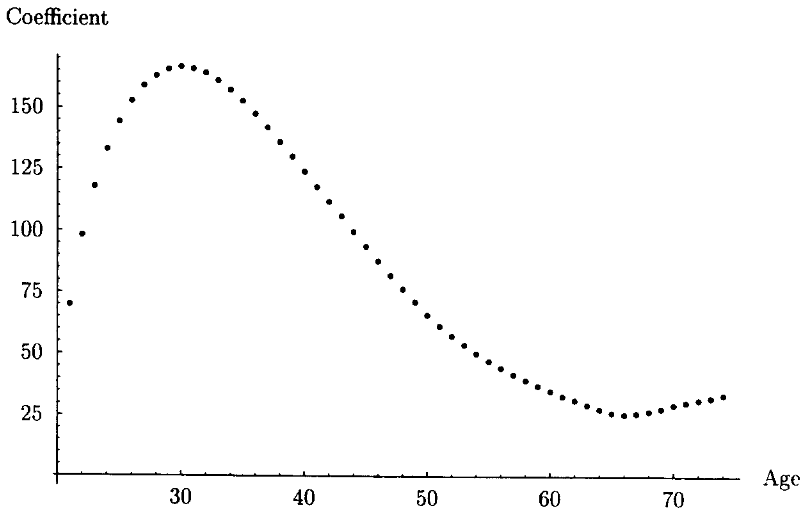

Fig. 2. Coefficients on σ_η^2 implied by HSZ parameters.

is clear from this figure that even if we were able to estimate age-specific coefficients they would not have been consistent with this model: between the ages of 25 and 50 the predicted coefficient on σ_η^2 from the HSZ model averages over 100, whereas our baseline empirical estimates of the coefficient on σ_η^2 for consumers in this age range varied only from 8.6 to 12.1, depending on the measure of wealth. Even if our estimated standard errors had been ten times larger, we would still have been able to reject a coefficient of 100 on the permanent variance term with an enormous degree of statistical significance. And the arguments for why our empirical estimates might have been biased downward (sample selection problems, pre-tax versus after-tax income variability) seem very unlikely (to us, at least) to have produced a coefficient estimate that was both highly significantly different from zero and roughly one-tenth the HSZ model's predicted value.

Deaton (1991) provides an insight into why the HSZ model implies such extraordinarily high responsiveness of wealth accumulation to the degree of uncertainty in permanent income. Deaton showed, using an infinite-horizon model where income was expected to grow at rate g indefinitely, that the successive consumption-to-permanent-income rules $c_T[x]$, $c_{T-1}[x]$, ... converge to a fixed rule $c^*[x]$ if the following 'impatience' condition is satisfied[22]

$$\rho^{-1}(r-\delta)+\frac{\rho}{2}\sigma_v^2<g. \tag{13}$$

[22] Carroll (1996) shows that the consumption rules converge in a model identical to the present one (i.e. without explicit liquidity constraints) under the same condition.

The Deaton condition amounts to an assumption that consumers are 'impatient' in the sense that, if there were no uncertainty, they would wish to spend more than their current income.

Consider what happens if the Deaton condition is not satisfied. In that case, consumers will always spend less than their income and therefore add to wealth. In this case, infinite-horizon consumers will accumulate wealth without bound. In the infinite-horizon model with uncertainty, if the Deaton condition is not satisfied the long-run elasticity of wealth with respect to the degree of permanent income uncertainty should be infinite. This provides the intuition for why the elasticities in the HSZ parameterization of the life cycle model are so high – their parametric choices come close to failing the Deaton condition. For their choices of $\rho = 3$, $r = \delta = 0.03$ and $\sigma_\eta = 0.10$, the left-hand side of Eq. (13) is $\rho^{-1}(r - \delta) = 0 + \rho/2\ \sigma_\eta^2 = 0 + 3/2\ (0.01) = 0.015$. It is a bit difficult to know the income growth rate for comparison, because in the finite-horizon context the income process cannot be summarized by a single growth rate. However, the average growth rate from age 25 to 65 under HSZ parameter values is close to 1.5% a year, and at age 65 income plunges to 70% of its pre-retirement level, so it is clear that their parametric choices are at least on the borderline of failing to satisfy the Deaton condition. Of course, in a finite-horizon context, wealth cannot be infinitely responsive to the degree of uncertainty as it is in the limit of an infinite-horizon model, but Fig. 2 shows that under the HSZ parameter values wealth can be remarkably sensitive to the degree of permanent uncertainty even in a finite-horizon model.

The foregoing argument suggests that the failure of the model under HSZ parameter values is a direct result of parametric assumptions which either fail, or come close to failing, the Deaton condition. However, under an alternative set of parameter values similar to those used in Carroll (1992, 1997), the model does indeed produce results that are much more consistent with our empirical findings.

5.3. Results using Carroll (1992, 1997) parameterization

The most important difference between Carroll (1992, 1997) parameter values and the HSZ parameter values is in the age–income profile. Using our PSID data we estimate empirical age-specific income growth rates that are uniformly from 1–2% points greater at every age than those used by HSZ; we estimate that income grows, on average, at roughly 3% a year from age 25 to age 55 and at 1% a year from age 55 to 65. This discrepancy arises because the HSZ age–income profiles were made from cross-section data that did not incorporate the portion of household income growth attributable to a generally rising level of productivity in the economy over time. We can see no valid reason to exclude this important component of income growth from the age–income profile.

Because the asset in the model is perfectly riskless and perfectly liquid, the closest real-world analogue to such an asset is probably short-term T-bills, whose post-war after-tax rate of return been close to 0% annually, so our baseline interest rate is 0. HSZ also assume a time preference rate of 3% annually; we follow much of the macroeconomics literature in choosing a time preference rate of 1% per quarter or 4% a year. A final difference is that HSZ do not model the possibility that there may be periods when income goes to zero. We adopt the Carroll (1992) estimate that zero-income events happen with annual probability 0.5%.

The solid line in Fig. 3 shows the results from solving the model under our alternative parameterization. Rather than growing steadily until retirement (as under the baseline HSZ parameter values), the net wealth ratio hovers around a constant 'target' level until around age 50. Only around age 50 does the wealth ratio begin rising sharply in anticipation of retirement. The characteristics of the kind of 'target' or 'buffer-stock' saving behavior that consumers in this model engage in before age 50 have been explored extensively in Carroll (1992, 1996, 1997). Those papers demonstrate that the logic behind the target-saving behavior is that precautionary motives prevent wealth from getting too low, while 'impatience' (in the Deaton sense described above) prevents assets from growing too large.

The dashed line in Fig. 3 shows the results when we again undertake the experiment of changing the value of the variance of permanent shocks from its baseline value of 0.1 to 0.05. The effect on the life-time profile of the net wealth

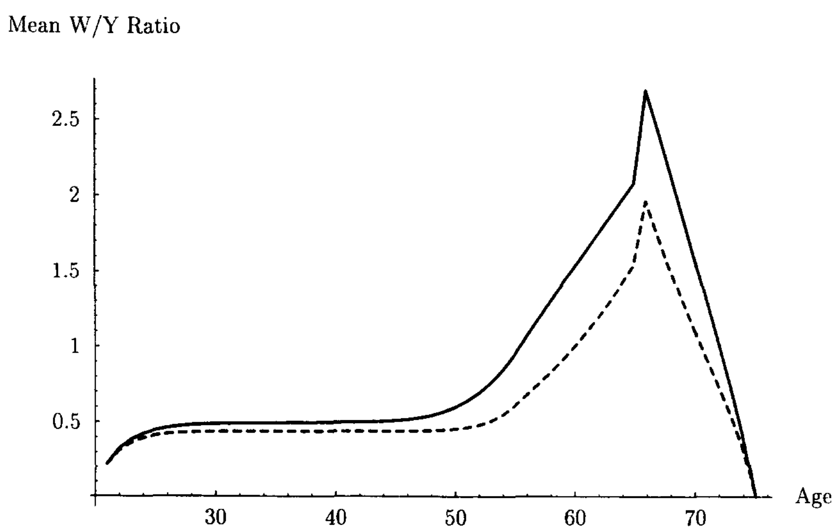

Fig. 3. Age–Wealth profile under Carroll parameters.

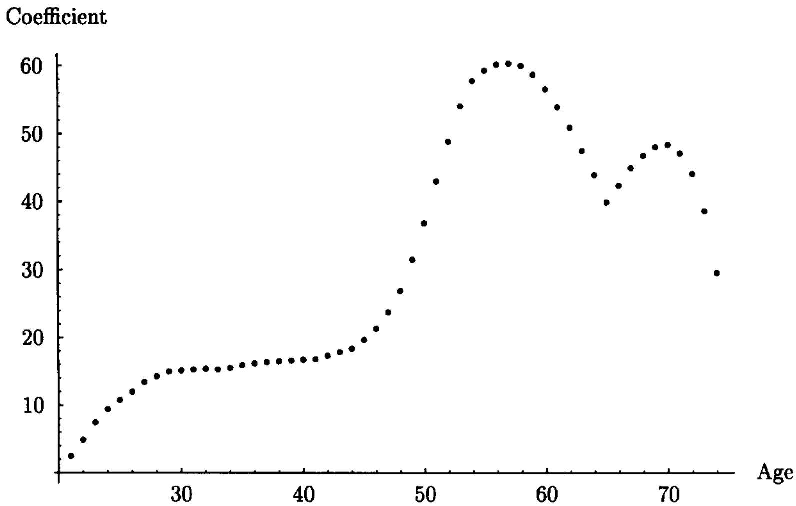

Fig. 4. Coefficients on σ_η^2 implied by Carroll parameters.

ratio is rather modest: the target wealth ratio declines by roughly the same amount at every age before 50. This impression is confirmed in Fig. 4, which calculates the life-time profile of the coefficient on σ_η^2. This figure is sharply different from the analogous Fig. 2 calculated using the HSZ parameter values, in two respects. First, rather than rising sharply in early life and then declining sharply later, under our parameter values the predicted coefficient is roughly constant from about age 25 to slightly before age 50. More importantly, however, the mean value of the coefficient over these ages is only about 20, far less than the values of over 100 that emerged under the HSZ parameter values. Although 20 is still above our typical empirical estimate of about 10, it is far closer than any of the predictions under HSZ parameter values. Note that, because all these parameter values are taken from Carroll (1992), there has been no direct calibration of the theoretical model to match the empirical result.

5.4. *Estimating consumers' time preference rates*

Further analysis will be easier if we turn our attention just to the consumers in this 'buffer-stock saving' phase of life, from roughly ages 25–50. Following Carroll (1992), we identify the 'target' level of wealth w^* as the level such that, if the consumer holds w^* in period t, then expected wealth in period $t+1$ is also w^*. For any given set of parameter values, we can solve for the target wealth that prevails during the 'buffer-stock' saving portion of the life cycle when income is growing at a baseline value of 3% a year. Our procedure for calculating the coefficient on uncertainty can then be formalized as follows. Write the log of target wealth as a function of all the parameters of the model, $\log(w^*) = f(\sigma_\eta^2, \sigma_\varepsilon^2, \rho, \pi, g, \delta)$ or, for notational simplicity, $\log(w^*) = f(\sigma_\eta^2, \boldsymbol{\theta}, \delta)$ where θ is the vector of other parameter values, $\boldsymbol{\theta} = \{\sigma_\varepsilon^2, \rho, \pi, g\}$. The coefficient of interest is the derivative of log wealth with respect to σ_η^2 around the baseline parameter values. Denoting the vector of baseline parameter values as $\bar{\theta}$ and the baseline value of the variance of the permanent shock as $\bar{\sigma}_\eta^2$, for a given value of the time preference rate $\delta = \bar{\delta}$ and for a small perturbation e around the baseline value of σ_η^2 we have

$$\left.\frac{\partial f(\sigma_\eta^2, \theta, \delta)}{\partial \sigma_\eta^2}\right|_{\theta = \bar{\theta}, \sigma_\eta^2 = \bar{\sigma}_\eta^2, \delta = \bar{\delta}} \approx \frac{f(\bar{\sigma}_\eta^2 + e, \bar{\theta}, \bar{\delta}) - f(\bar{\sigma}_\eta^2 - e, \bar{\theta}, \bar{\delta})}{2e}. \tag{14}$$

This equation highlights the fact that, if f is nonlinear, the coefficient on the variance of permanent shocks will in general depend upon our assumptions about the baseline values for all of the parameters of the model. With the exception of the time preference rate, δ, our baseline parametric choices are consistent with empirical estimates in the literature. In our view, however, there is little credible evidence on the appropriate value of δ.[23] One way to interpret our empirical results, therefore, is by asking what values of δ would be consistent with our empirical estimates. That is, taking as fixed the values of the other parameters, for a given estimate of the coefficient on σ_η^2 (call it C), we can search for the δ such that

$$\left.\frac{\partial f(\sigma_\eta^2, \theta, \delta)}{\partial \sigma_\eta^2}\right|_{\theta = \bar{\theta}, \sigma_\eta^2 = \bar{\sigma}_\eta^2} \approx C. \tag{15}$$

This idea is illustrated in Fig. 5, which shows the results of performing the calculation on the right hand side of Eq. (14) for a wide range of possible values for δ, under the baseline values specified above for the other parameters. This should roughly correspond to the model's prediction, for each value of δ, of the coefficient in a regression like our empirical specification (11).

For the baseline specification in Table 2 with NHNBW as the dependent variable, the estimate of the coefficient on σ_η^2 was 9.4, with a standard error of 3.0. As shown in the figure, the value of δ that corresponds to an σ_η^2 coefficient of 9.4 is 10.6% per year; thus, our empirical estimate of the coefficient on NHNBW in Table 2 implies a point estimate of the rate of time preference of 10.6% per year. Unfortunately, however, the two-standard-deviation bands encompass a very wide range of possible values of δ; the lower bound is 4.7% and the upper nearly 50%.

[23] One of the few papers to attempt an empirical estimate of δ is Lawrance (1991). The results of that study are critiqued in Carroll (1997).

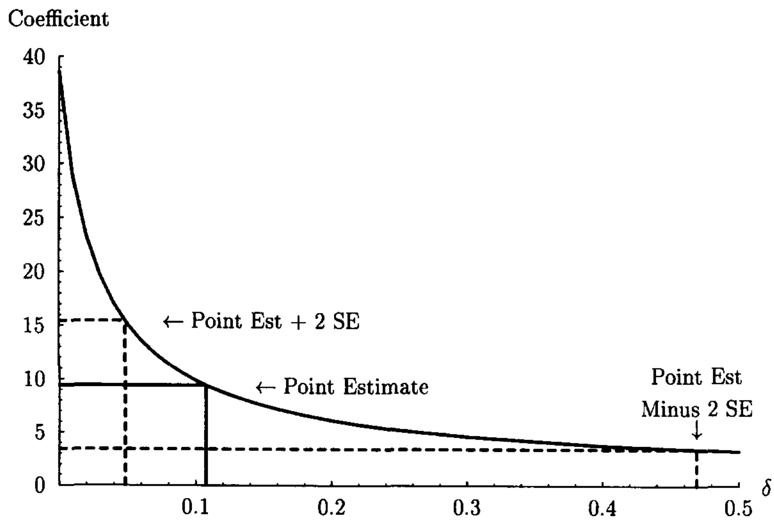

Fig. 5. Estimated time preference rate under buffer stock model.

Because there is room for disagreement with our choices for baseline values of other parameters, Table 4 shows the point estimate and the two-standard-error range for δ if we try two alternative values for each of the other important parameters in the model. Unfortunately, the conclusion about the location of δ is rather sensitive to the choices of other parameter values, with the point estimates ranging from about 5% to about 14% per year.

Of course, if we choose an entirely different constellation of parameter values, rather than just allowing them to diverge one-by-one from our baseline assumptions, the implied estimate of δ can change more. This is illustrated by the last row in the table, which reflects the results when all parameter values except δ are set to HSZ values and income growth is assumed to proceed at 2% a year (which corresponds to the most rapid rate of income growth estimated by HSZ for any period of life). The point estimate of the time preference rate is 38%, and the two-standard-error band ranges from 21% to 79%. (If income growth had been assumed to be lower, the estimates of δ would have been even higher.) Thus, the model solved under HSZ values for all parameters other than δ is consistent with our empirical results only if the time preference rate is at least 21% annually, rather than the 3% that HSZ assumed. Such a high time preference rate strikes us as implausible.

The reason that consumers engaged in buffer-stock saving behavior do not react strongly to uncertainty in permanent income is that, speaking in the loose sense of Friedman (1957), these consumers act as though they have a short 'horizon'. More formally, Carroll (1997) shows that, under the same 'buffer-stock'

Table 4
Discount rates implied by baseline parameterizations of intertemporal consumption problem, mean and two standard error confidence intervals

Parameters	Point estimate of discount rate	Two-standard errors below	Two-standard errors above
Baseline parameters	0.107	0.048	0.469
Interest rate = 0.02	0.130	0.069	0.496
Interest rate = 0.04	0.152	0.079	0.540
Growth rate = 0.02	0.140	0.079	0.513
Growth rate = 0.04	0.075	0.018	0.425
Coefficient of relative risk aversion = 2	0.084	0.034	0.387
Coefficient of relative risk aversion = 4	0.134	0.067	0.552
Standard deviation of permanent shock = 0.05	0.046	− 0.001	0.331
Standard deviation of permanent shock = 0.15	0.183	0.118	0.611
Standard deviation of transitory shock = 0.05	0.102	0.046	0.445
Standard deviation of transitory shock = 0.15	0.115	0.052	0.501
Probability of zero-income event = 1%	0.096	0.040	0.432
Probability of zero-income event = 0.1%	0.136	0.067	0.551
Hubbard–Skinner–Zeldes parameters	0.383	0.210	0.786

Note: See Section 5 in the text for discussion.

parameter values used in this, consumers holding the target amount of wealth will behave, at least in some circumstances, as though they discount future income at an average rate of around 22%,[24] implying a Friedman-style horizon of less than 5 years. The reason why impatient buffer-stock consumers discount future income so heavily is precautionary; they are unwilling to spend today a dollar that in expectation will, but just possibly might not, arrive tomorrow. By contrast, for consumers who are as patient as HSZ assume, future income is discounted very little (close to the real interest rate), and so a risk to the entire stream of future income (i.e. the risk of a permanent shock) is much more dramatic than for short-horizon consumers.

6. Conclusion

This paper presents some of the first empirical evidence that consumers who face greater income uncertainty hold more wealth. We decompose income

[24] See Carroll (1997) for the exact experiment.

uncertainty into a variance of transitory shocks and a variance of permanent shocks, and show that both transitory and permanent uncertainty are statistically significant in predicting levels of household wealth for three different definitions of wealth.

We further show that our empirical results are inconsistent with a parameterization of the life cycle model advocated by Hubbard et al. (1994) in which the time preference rate is low enough for retirement saving to be important at early ages. A model with such patient consumers predicts a very large response of wealth holdings to the degree of uncertainty in permanent income – the predicted regression coefficient on the variance of permanent income is roughly ten times larger than the coefficient we estimated. These extreme results are not due to unusual assumptions about consumers' prudence (the intensity of their precautionary saving motive) or about the magnitude of income shocks (which, if anything, are conservatively chosen in the simulations). Rather, they arise because rational consumers realize that permanent shocks to income will last the rest of their lives; if consumers are patient and even modestly prudent, the present discounted utility effect from a possible negative shock to permanent income is very large and therefore justifies a large saving response.

The inability of the life cycle model under the HSZ parameterization to explain our empirical results does not necessarily imply that consumers are irrational. Instead, we show that the empirical results are consistent with a 'buffer stock' saving configuration of the model that emerges when consumers are more impatient than in the HSZ parameterization and face an income growth path more consistent with experience in US data. In the buffer stock version of the model, rather than saving for retirement from very early in life, consumers spend most of their working life-times trying to maintain a modest 'target' wealth–income ratio; they begin saving for retirement only around age 50.

Appendix

A.1. Sample restrictions imposed

We restrict the sample to households that remained intact over the chosen sample period. According to our definition, a household remains intact whenever the same person is the head of household each year and the spouse, if present, remains the same. We also exclude those households whose income in any year was less than 20% of its average over the period. This exclusion is necessary in order to calculate our measures of uncertainty; if these households are included our results tend to be almost entirely dominated by a few observations. We also restrict the sample by excluding those households who were included in the PSID's poverty sample. Finally, we use income data only from the years 1981–87, with the idea that choosing a period centered around 1984 will give us the best possible estimates of the degree of income uncertainty near that year. Results were similar when the sample was extended to include the years 1976–87.

A.2. Estimating permanent and transitory variances in the sample

A.2.1. The projection methodology

As described in the text, the variance of an income difference of length d is given by (Eq. (6), Section 3):

$$\mathrm{Var}(r_{\mathrm{d}}) = d \cdot \sigma_\eta^2 + 2 \cdot \sigma_\varepsilon^2 \tag{A.1}$$

where σ_η^2 and σ_ε^2 are the variances of the permanent and transitory shocks to income, respectively. We estimate $\mathrm{Var}(r_{\mathrm{d}})$ for each household by

$$v_{\mathrm{d}} = r_{\mathrm{d}}^2 = \mathrm{Var}(r_{\mathrm{d}}) + \mu_{\mathrm{d}}, \tag{A.2}$$

where μ_{d} is a mean-zero disturbance. Using the previous equation to substitute for $\mathrm{Var}(r_{\mathrm{d}})$ yields the following regression equation:

$$v_{\mathrm{d}} = d \cdot \sigma_\eta^2 + 2 \cdot \sigma_\varepsilon^2 + \mu_{\mathrm{d}}, \tag{A.3}$$

where observations are distinguished by the length of the difference d. Our method simply does OLS household by household of $\boldsymbol{v} = \{v_{\mathrm{d}}(1), \ldots, v_{\mathrm{d}}(n)\}'$ on $[\boldsymbol{d} \quad \mathbf{2}]$, where $\boldsymbol{d} = \{d(1), \ldots, d(n)\}'$, $\mathbf{2} = \{2, \ldots, 2\}'$, and $\mathrm{d}(t)$ is the tth household difference. As discussed below, we use $n = 9$ for each household. The coefficients obtained for this regression give household estimates of σ_η^2 and σ_ε^2, which we have denoted s_η^2 and s_ε^2 in the text.

We have assumed that μ_{d} and $\mu_{\mathrm{d}'}$ are i.i.d. for each household, making OLS the efficient way to conduct the estimation. There does not appear to be any reason to believe that the noise in observing r_{d}s should vary with d for a given observation, as the data is collected each year and r_{d} represents a *difference* in annual incomes.

A.2.2. Robustness to MA(q) serial correlation

The variance decomposition methodology we have adopted can be made robust to MA(q) serial correlation in the transitory shock. This can be seen in Eq. (5) from Section 3:

$$r_{\mathrm{d}} = \{\eta_{t+1} + \eta_{t+2} + \ \ldots \ + \eta_{t+d}\} + \varepsilon_{t+d} - \varepsilon_t, \tag{A.4}$$

by noting that $\mathrm{Var}(\varepsilon_{t+\mathrm{d}} - \varepsilon_t) = 2{*}\sigma_\varepsilon^2 - 2{*}\mathrm{Cov}(\varepsilon_{t+\mathrm{d}}, \varepsilon_t)$. This will yield an unbiased estimate of σ_ε^2 whenever $\mathrm{Cov}(\varepsilon_{t+\mathrm{d}}, \varepsilon_t) = 0$. Thus, as long as we restrict our choices of d to those greater than q, the procedure will yield the unbiased estimate. Results from MaCurdy (1982) and Abowd and Card (1989) suggest that there is no evidence of serial correlation beyond order 2 for labor income of the household head.

We choose all possible pairs of income values in which the years are at least three apart (thus making our estimates robust to serial correlation up to MA(2)) and for which more than one pair of years is available to estimate that difference. This leaves us with nine differences for the period 1981–87:

Length (d)	Years used
5	1981–86, 1982–87
4	1981–85, 1982–86, 1983–87
3	1981–84, 1982–85, 1983–86, 1984–87

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