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Does the choice of consumption measure matter? An application to the permanent-income hypothesis

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Abstract

Food consumption in the Panel Study of Income Dynamics is used as a proxy for total consumption in many applications in economics, including tests of the permanent-income hypothesis, tests of separability between consumption and leisure, and tests of intergenerational altruism. Food, however, explains only a small fraction of the variation in total consumption. I propose a measure of composite consumption based on predicted wealth and compare it both to food consumption and to Skinner's (1987) measure of predicted consumption in a test of the permanent-income hypothesis. Using a log-linear intertemporal consumption function I find that food does not reject the permanent-income hypothesis but both Skinner's predicted consumption and the composite measure proposed here do reject the hypothesis. © 1998 Elsevier Science B.V. All rights reserved.

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1. Introduction

Food consumption is the primary consumption measure collected in the Panel Study of Income Dynamics (PSID), and is used as a proxy for total

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consumption in many applications, including tests of the permanent-income hypothesis (Hall and Mishkin, 1982; Shapiro, 1984; Altonji and Siow, 1987; Zeldes, 1989; Runkle, 1991; Mariger and Shaw, 1993), tests for separability between household consumption and leisure choices (Altonji, 1986; Shaw, 1989; Altug and Miller, 1990), and tests for intergenerational altruism (Altonji et al., 1992). The consumption of food, however, only accounts for 39% of the variation in total consumption and thus may lead to biased tests if it is a poor approximation of total consumption (Skinner, 1987). In the current paper I propose a new measure of composite consumption using predicted wealth calculated from the PSID and examine the sensitivity of tests of the permanent-income hypothesis to the choice of consumption.

Recently, the question of whether consumers obey the central prediction of the permanent-income hypothesis, that consumers base current consumption decisions on expected lifetime income rather than current income, has shifted from aggregate time-series tests (Hall, 1978; Flavin, 1981; Mankiw, 1981) to household-based tests from longitudinal data (Hall and Mishkin, 1982; Shapiro, 1984; Hayashi, 1985; Altonji and Siow, 1987; Zeldes, 1989; Runkle, 1991; Mariger and Shaw, 1993; Attanasio and Weber, 1995; Lusardi, 1996). Household-based tests of the hypothesis are superior to aggregate time-series tests because they permit identification of consumers who are potentially liquidity constrained. However, evidence of liquidity constraints in household analyses is mixed (Deaton, 1992; Browning and Lusardi, 1996).

The discrepancy in tests of whether consumption is 'excessively sensitive' to income changes may partially be explained by the use of food consumption, which is likely an inappropriate proxy for total consumption. Specifically, if individuals experience a temporary negative shock to income, then food consumption can be stabilized with food stamps. Moreover, the income elasticity of food demand is, in general, smaller than other components of consumption; hence, excess-sensitivity tests based on food are likely to lack power. For example, Blundell et al. (1993) report an income elasticity of food demand of 0.61; whereas, the income elasticities of demand for other consumption components are 2.29 for alcohol, 0.92 for clothing, 1.45 for services, and 1.20 for transport. In addition, the use of food consumption implicitly assumes that utility is additively separable from nonfood consumption. If the additive separability assumption is incorrect, as suggested in Attanasio and Weber (1995), then the Euler equation on which the test is based may be invalid. While most authors have dealt with the problem of distinguishing between expenditures on durable goods as opposed to the service flow of durables by using nondurable consumption, it is possible that the avoidance of durable consumption has been carried to an extreme in the case of food, leading to the potentially more serious problems mentioned above. Consequently, it is important to conduct tests using broad measures of consumption to examine if the results are robust.[1]

I examine the issue of whether tests of the permanent-income hypothesis are sensitive to the choice of consumption measure using Runkle's (1991) log-linear approximation to the intertemporal consumption decision. Runkle's framework is employed because of his careful attention to problems induced by measurement errors-in-variables, persistence in household consumption, and aggregate shocks. Relying on food consumption for the years 1973–1982 from the PSID, he failed to reject the permanent-income hypothesis for the whole sample, for subsets of homeowners and renters, and for households with more than or less than two months' income in liquid assets. I confirm Runkle's findings with my PSID sample for the years 1977–1986 using food consumption, but reject the permanent-income hypothesis using both the composite measure proposed here and a measure of predicted consumption proposed by Skinner (1987). Moreover, I show that Zeldes's rejection of the permanent-income hypothesis using food consumption is likely due to his use of an inconsistent estimator and not due to different sample selection criteria relative to Runkle (1991) or this paper.

2. The consumption measures

The PSID is the longest and most comprehensive longitudinal data set in the United States with socio-economic information collected on over 37,000 individuals between the years 1968 and 1989 (Hill, 1992). Over the years, information has been gathered on various components of consumption such as food consumed at home and at restaurants, the net value of food stamps, annual housing payments for owners and renters, annual utility payments, and the number of vehicles owned. The value of personal savings was collected in each of the first five waves but was discontinued thereafter. Most researchers using the PSID data rely on food for consumption- based analyses because it has the

[1] Hayashi (1985), Attanasio and Weber (1995), and Lusardi (1996) are similar to this study in their use of broad consumption measures. Hayashi models durable-goods expenditures in an application to Japanese panel data, finding evidence of excess sensitivity. Attanasio and Weber examine nondurable consumption using cohort averages over time from the Consumer Expenditure Survey (CEX), finding no evidence of excess sensitivity after accounting for within-period nonseparability between consumption and labor supply. However, the consistency of their estimates hinges on the assumption that the preferences of individuals are accurately represented by the average preferences of the assigned cohort. Moreover, it is difficult to interpret their results as a nonrejection of the hypothesis because hours changes are a good proxy for income changes such that the inclusion of hours may simply be capturing the influence of earnings. Lusardi finds evidence of excess sensitivity using nondurable consumption from the CEX but income data from the PSID. Her results, however, are limited in that only one consumption change is observed for each household.

longest time-series coverage. For the present purpose, and following Runkle, food consumption in time $t(FC_t)$ is defined as the sum of food consumed at home, food consumed in restaurants, and the net value of food stamps. Real values are derived by deflating the nominal food consumption with the food component of the consumer price index.

Skinner (1987) suggested a broader measure of predicted consumption as an alternative to food in the PSID. The components in Skinner's measure include food consumption, house value, rent payments, utility payments, and the number of vehicles owned. Rent payments for homeowners are computed as the imputed rental value of the home, assumed to be 6% of its market value. Using data from the Consumer Expenditure Survey in 1972–1973 and 1983, Skinner estimated that the latter components explain 78% of the variation in total consumption in 1972–1973 and 73% of the variation in 1983. Moreover, he demonstrated that the coefficients are relatively stable over time; hence, he recommended using the 1972–1973 coefficients in constructing predicted annual consumption. The specification of predicted consumption (PC_t) is (t-ratios in parentheses)

$$\begin{aligned} PC_t = \underset{(2.4)}{110.1} &+ \underset{(86.8)}{1.418}*\mathit{food(home)} + \underset{(86.0)}{2.604}*\mathit{food(away)} + \underset{(80.8)}{0.0988}*\mathit{house\ value} \\ &+ \underset{(87.9)}{1.538}*\mathit{rent} + \underset{(34.1)}{2.257}*\mathit{utilities} + \underset{(26.0),}{624.6}*\mathit{vehicles}, \\ &\qquad N = 14{,}499 \qquad \bar{R}^2 = 0.777, \end{aligned} \tag{1}$$

where food is deflated by the food component of the CPI while the other expenditures are deflated by the annual personal consumption expenditure deflator to give real predicted consumption.

I propose a new alternative to food consumption in the PSID based on household wealth (A_t), defined as the sum of liquid assets and home equity. In particular, liquid assets are constructed by dividing the first \$200 of rent, interest, and dividend income by the average annual passbook savings rate and the remaining income by the average annual 3-month T-bill rate (Zeldes, 1989; Runkle, 1991). Home equity is the difference between the annual market value of the home and the remaining mortgage principal. Given wealth, personal saving is found by taking the year-to- year change in wealth; that is, $S_t = A_{t+1} - A_t$. Composite consumption follows directly by subtracting saving from disposable personal income, $CC_t = Y_t - S_t$. The composite measure is advantageous over predicted consumption because Skinner's measure may be unstable if the relative prices of goods change, as documented in Attanasio and Weber (1995). Moreover, many of the components needed to construct his measure are no longer collected by the PSID.

Table 1
OLS and first-difference wealth regressions for 1984 and 1989[a]

	Wealth 1984		Wealth 1989		First differences 1984–1989	
Constant	43875.160 (2976.81)	9891.524 (2046.448)	66179.719 (3974.91)	15410.406 (2966.51)	—	—
Liquid assets	1.798 (0.29)	1.237 (0.287)	1.155 (0.23)	0.683 (0.14)	0.599 (0.281)	0.418 (0.190)
Home equity		1.200 (0.062)		1.357 (0.07)		1.299 (0.121)
$\bar{R}^2$	0.599	0.783	0.526	0.812	0.837	0.902

[a] $N = 1116$. Heteroskedasticity-corrected standard errors are in parentheses.

A key issue in constructing CC_t is how well the liquid assets and home equity approximate total household wealth. The PSID collected comprehensive data on household wealth in the 1984 and 1989 waves, permitting such an examination. The net wealth data in 1984 and 1989 include the sum of house value, net value of other real estate, net value of vehicles, net value of farm or business, net value of stocks, value of cash accounts, and the net value of other assets less remaining mortgage principal and other debts. The first four columns of Table 1 contain the results of regressing total wealth on liquid assets and a constant, with and without home equity, for 1984 and 1989.[2] Including home equity improves the fit of the model substantially, with the adjusted R^2 increasing from 0.599 to 0.783 in 1984 and from 0.526 to 0.812 in 1989. On average, a constant, home equity, and liquid assets explain about 80% of the variation in total wealth.

Because we observe the same individual's wealth holdings in two periods it may be possible to improve the fit of predicted total wealth by pooling the data and permitting person-specific heterogeneity in the intercepts. Before pooling, though, it is necessary to establish that the estimated coefficients are structurally stable across the periods. The standard Chow test is inappropriate for testing structural change in the presence of heteroskedasticity; however, the Wald test is

[2] The sample selected for the regressions in Table 1 consists of heads of households who do not change marital status (either married or single continuously) between 1976 and 1989, are not self-employed, and do not have missing data on any of the consumption, income, or tax variables resulting in a balanced panel of 1116 households for 14 years, or 15,624 person-years.

heteroskedasticity-robust and is given as

$$(\hat{\theta}_{84} - \hat{\theta}_{89})'[V(\hat{\theta}_{84}) + V(\hat{\theta}_{89})]^{-1}(\hat{\theta}_{84} - \hat{\theta}_{89}) \sim \chi^2(k), \tag{2}$$

where $\hat{\theta}_t$ ($t = 1984, 1989$) is the ($k \times 1$) vector of estimated parameters and $V(\hat{\theta}_t)$ is the ($k \times k$) covariance matrix. Under the null hypothesis of stability the value of the test statistic comparing the model with both liquid assets and home equity as regressors is 11.13 with a p-value of 0.011 at 3 degrees of freedom. While the Wald test offers some support for the null hypothesis of structural stability, an additional test of stability is to determine whether the estimated 1984 coefficients are effective in predicting 1989 wealth. For this test I regress predicted 1989 wealth on predicted 1984 wealth using 1989 data, resulting in a coefficient of 0.90 and an adjusted R^2 of 0.96. Hence, I conclude that the estimated coefficients are relatively stable and the data can be pooled for a first-difference regression.[3]

In the last two columns of Table 1 I report the results from first-difference regressions of total wealth on liquid assets (col. 5) and on liquid assets and home equity (col. 6). The fit of the model improves markedly; the adjusted R^2, constructed for the levels of wealth with person- specific intercepts included, increases to 0.84 in the model with liquid assets and to 0.90 with home equity included. Because the latter specification provides the best prediction for total wealth it will also give the best prediction for changes in wealth used to construct composite consumption. Consequently, I use the parameters in column (6), along with the estimated person-specific intercepts, to construct composite consumption in the application below.[4]

I now turn to a brief graphical comparison of the three consumption measures. I focus on the years 1977–1986 because it is the only period where there is simultaneous collection of food, vehicles, and utilities. In order to minimize measurement error, observations are deleted if consumption was less than or equal to zero, grew by more than 200% in a year or fell by more than 90% in a year. The ensuing unbalanced 10-year panel has 8629 person-years.

Carroll and Summers (1991) used a simple graphical exposition to show that consumption 'tracks' income too closely to be consistent with the permanent-income hypothesis. It is instructive to examine whether the life-cycle path of

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[3] The advantage of a first-difference estimator over the standard 'within' estimator is that the former can eliminate possible nonstationarities in wealth; however, in this application, the estimated coefficients coincide across both transformations.

[4] In an earlier version of the paper I used the parameters in column (2) to construct composite consumption with no qualitative difference in the results. While the fixed-effect model gives the best overall fit, a limitation of these coefficients is that they are sample-specific and not generalizable to the whole population. This problem plagues inference in all fixed-effects models (Hsiao, 1986). Random effects coefficients are generalizable to the whole population, but the fit of the random effects model was much weaker with an adjusted R^2 of 0.73.

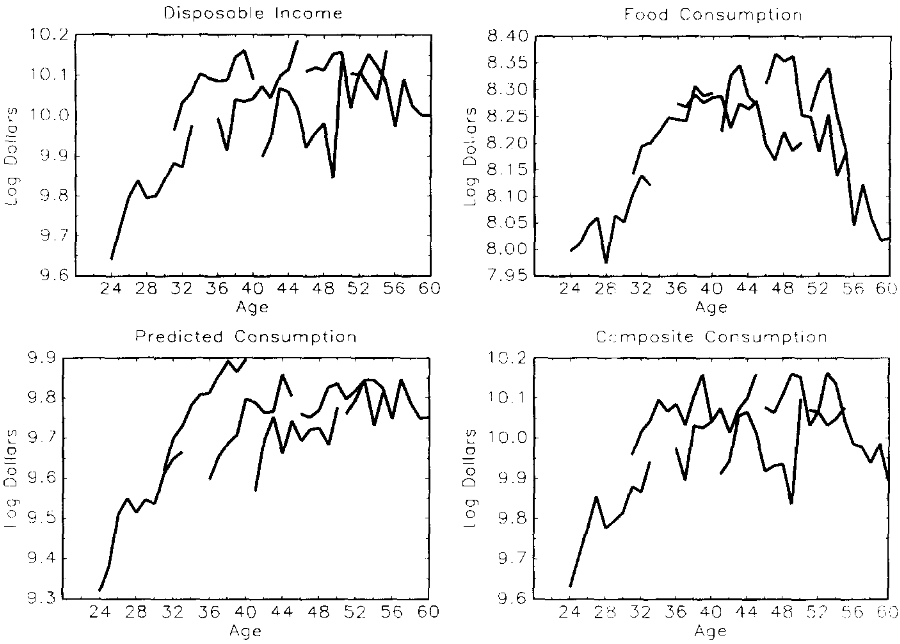

Fig. 1. Consumption and income over the life cycle.

food consumption is more consistent with the hypothesis relative to the broader consumption measures. To this end, I construct six birth- year cohorts, each of which span 10 years, and regress each consumption measure and disposable income on 60 cohort-year dummies. The coefficients on the cohort-year dummies, representing cohort-year averages, are plotted in Fig. 1. It is clear from the diagrams that both predicted consumption and composite consumption track disposable income 'more closely' than food consumption, especially after age 46; however, food consumption is less smooth than suggested by the permanent-income hypothesis and formal econometric tests are necessary to make more definitive claims on the 'excess sensitivity' of consumption to income changes.

3. Testing the permanent-income hypothesis

The framework for testing the permanent-income hypothesis is found in Runkle (1991) and the interested reader is referred there for further details. The consumer is assumed to maximize the present discounted value of uncertain, time-separable utility subject to an asset accumulation constraint. Furthermore, contemporaneous consumption and leisure choices are additively separable and the instantaneous utility function is parameterized with the constant relative risk-aversion utility function. Because of potential measurement error in consumption and the ensuing difficulty of estimating nonlinear errors-in-variables models, the intertemporal consumption decision is approximated with a log-linear variant given as

$$\Delta C_{i,t+1} = \ln(C_{i,t+1}) - \ln(C_{i,t}) = \alpha_0 + \alpha_1 r_{i,t} + \varepsilon_{i,t+1}, \tag{3}$$

where $\ln(\cdot)$ is the natural-log operator, α_0 is a constant term and is a function of the subjective discount rate, $\alpha_1 = (1+\alpha)^{-1}$ is the intertemporal substitution elasticity measuring the percentage change in consumption growth due to a 1% change in the interest rate, $r_{i,t}$ is the real after-tax interest rate, and $\varepsilon_{i,t+1}$ is a forecast error assumed independent of all information available in time t, $E(\varepsilon_{i,t+1}|I_{it}) = 0$. If measurement error exists in consumption, then $\varepsilon_{i,t+1} = v_{i,t+1} + \mu_{i,t+1} - \mu_{i,t}$, where $v_{i,t+1}$ is a random error and $\mu_{i,j}\,(j = t+1, t)$ is measurement error. In this case the forecast error is independent of predetermined information in time $t-1$.

3.1. Estimation issues

Because of the potential correlation between the explanatory variables and the forecast error arising from measurement error, Hansen (1982) generalized method-of-moments (GMM) estimator is used to produce consistent and efficient estimates of Eq. (3). The GMM estimator minimizes the optimally weighted criterion function

$$J_T = g(\alpha)' W g(\alpha), \tag{4}$$

where $g(\alpha)$, $\alpha = [\alpha_0, \alpha_1]'$, is the sample average of the orthogonality conditions implied by Eq. (3) and W is an optimal weight matrix that permits conditional heteroskedasticity and autocorrelation. Initial consistent estimates for ε_{t+1} used in constructing W are obtained by setting the weight matrix to the identity matrix and estimating the model with two-stage least squares. The sample value of the criterion function, $\hat{J}_T$, is asymptotically distributed chi-square with L degrees of freedom, where L is the number of overidentifying restrictions imposed in estimation. Consequently, the J-statistic can be used as a model specification test of the validity of the overidentifying restrictions.

In selecting the correct specification to test the permanent-income hypothesis with food consumption, Runkle tested for the presence of persistent household-specific differences in consumption growth, whether an MA(1) specification for measurement error is sufficient to capture the time-series properties in consumption growth, and if there are aggregate shocks to consumption growth not captured by the after-tax interest rate. I conduct a similar model specification procedure for each of the consumption measures and concur with Runkle's specification for both food consumption (FC_t) and predicted consumption (PC_t). However, measurement error seems to be more of a problem with composite consumption (CC_t) and I conclude that a MA(2) error structure is more appropriate in this case.

3.2. Results from the permanent-income tests

The permanent-income hypothesis states that as long as the instrument set consists only of information known to the household prior to time t then no other explanatory variable should significantly determine consumption growth in Eq. (3). The instrument set includes a constant term and the $(t-1)$ and $(t-2)$ values of the householder's age, disposable personal income (defined as total family money less federal and social security taxes paid), the head's annual hours of work, the value of asset income, the value of liquid assets, and the after-tax real passbook and T-bill interest rates. Following Runkle, tests for excess sensitivity in consumption growth are conducted by appending Eq. (3) with each of the four additional regressors: $\ln Y_{i,t}$, $\ln Y_{i,t-1}$, $\Delta \ln Y_{i,t} = \ln Y_{i,t} - \ln Y_{i,t-1}$, and $\Delta \ln Y_{i,t-1} = \ln Y_{i,t-1} - \ln Y_{i,t-2}$.

3.2.1. Food consumption

Table 2 contains the regression results for food consumption using the entire 1977–1986 sample. All regressions include the head's age as a regressor to control for preference heterogeneity. The base case of Eq. (3) is in column (1) where the intertemporal substitution elasticity is estimated to be 0.22 with a standard error of 0.077. The J-statistic testing the validity of the 13

Table 2
Tests of the permanent-income hypothesis using food consumption: GMM estimator[a]

	(1)	(2)	(3)	(4)	(5)
Constant	0.0835	−0.1419	0.1367	0.0842	0.0838
	(0.0120)	(0.0448)	(0.0429)	(0.0124)	(0.0121)
After-tax interest rate	0.2191	0.4253	0.4118	0.2347	0.2069
	(0.0770)	(0.1746)	(0.1717)	(0.1070)	(0.0850)
Extra regressor		−0.0050	−0.0045	0.0081	−0.0064
		(0.0036)	(0.0034)	(0.0414)	(0.0266)
		$\ln Y_{\cdot,t}$	$\ln Y_{\cdot,t-1}$	$\Delta \ln Y_{\cdot,t}$	$\Delta \ln Y_{\cdot,t-1}$
J-statistic	20.9334	19.3188	19.4415	20.8929	20.890
	[13, 0.074]	[12, 0.081]	[12, 0.078]	[12, 0.052]	[12, 0.052]

[a] $N = 4261$. Standard errors are in parentheses while the degrees of freedom and p-value for the overidentifying-restrictions test are reported in square brackets. All GMM regressions have a MA(1) weight matrix.

overidentifying restrictions is 20.93, which has a p-value of 0.074 from the χ^2 distribution. As a comparison, Runkle, who uses a sample from the PSID from 1973–1982, finds an intertemporal substitution elasticity of 0.45 with standard eror of 0.16. More importantly, columns (2)–(5) suggest that food consumption growth does not exhibit excess sensitivity to income changes because the income variable is not statistically significant. The results in Table 2 concur with Runkle that the permanent-income hypothesis is not rejected for food consumption.

Some authors have suggested that Zeldes's (1989) rejection of the hypothesis using food consumption and Runkle's (1991) nonrejection is possibly due to differences in sample selection (Deaton, 1992; Browning and Lusardi, 1996). Specifically, Zeldes selects a much broader sample in the cross-sectional dimension than either Runkle or myself by his inclusion of split-off households, such that our narrower samples may be suppressing important heterogeneity in the data that leads to a rejection of the hypothesis. Rather than reconstructing Zeldes's sample, I indirectly address this issue by re-estimating the consumption growth equation for food using Zeldes's two-stage 'within' estimator; i.e., Zeldes assumed that there exist permanent, time-invariant differences in consumption growth across households, which he swept away by using the within transformation. This estimator, as noted by Keane and Runkle (1992), is inconsistent in the context of a rational-expectations model where the only instruments available are predetermined. In Table 3 I present two-stage within estimates of food consumption growth. Focusing on Zeldes's preferred specification (column (2)) yields a significant coefficient on the income term of -0.158 (0.0405). Hence, Zeldes's rejection of the permanent-income hypothesis using food consumption is likely due to his use of the inconsistent within estimator

Table 3
Tests of the permanent-income hypothesis using food consumption: within estimator[a]

	(1)	(2)	(3)	(4)	(5)
Constant	—	—	—	—	—
After-tax interest rate	1.4774	0.9742	1.3896	1.4110	1.2851
	(1.1106)	(1.1024)	(1.1069)	(1.1104)	(1.1069)
Extra regressor		-0.1584	-0.0382	-0.0759	-0.0627
		(0.0405)	(0.0366)	(0.0285)	(0.0454)
		$\ln Y_{\cdot,t}$	$\ln Y_{\cdot,t-1}$	$\Delta \ln Y_{\cdot,t}$	$\Delta \ln Y_{\cdot,t-1}$
J-statistic	56.3018	39.9913	50.5863	69.4567	42.5346
	[12, 0.000]	[11, 0.000]	[11, 0.000]	[11, 0.000]	[11, 0.000]

[a] $N = 4261$. Standard errors are in parentheses while the degrees of freedom and p-value for the overidentifying-restrictions test are reported in square brackets. All within regressions have a MA(1) weight matrix.

(i.e. note the rejection of the overidentifying test) and not due to a different sample selection.

3.2.2. Predicted and composite consumption

Tables 4 and 5 contain GMM regressions parallel to Table 2 for predicted consumption and composite consumption. The conclusions, however, are substantially different. The permanent-income hypothesis is rejected in Table 4

Table 4
Tests of the permanent-income hypothesis using predicted consumption[a]

	(1)	(2)	(3)	(4)	(5)
Constant	0.0527	0.2164	0.2080	0.0511	0.0547
	(0.0108)	(0.0544)	(0.0512)	(0.0105)	(0.0113)
After-tax interest rate	− 0.1872	0.0795	0.0732	− 0.1701	− 0.1860
	(0.0948)	(0.1296)	(0.1281)	(0.0928)	(0.0937)
Extra regressor		− 0.0152	− 0.0145	0.0790	− 0.0196
		(0.0050)	(0.0046)	(0.0492)	(0.0246)
		$\ln Y_{\cdot,t}$	$\ln Y_{\cdot,t-1}$	$\Delta \ln Y_{\cdot,t}$	$\Delta \ln Y_{\cdot,t-1}$
J-statistic	20.9284	11.6873	11.8267	19.7350	20.376
	[13, 0.0744]	[12, 0.471]	[12, 0.459]	[12, 0.072]	[12, 0.060]

[a]$N = 4261$. Standard errors are in parentheses while the degrees of freedom and p-value for the overidentifying-restrictions test are reported in square brackets. All GMM regressions have a MA(1) weight matrix.

Table 5
Tests of the permanent-income hypothesis using composite consumption[a]

	(1)	(2)	(3)	(4)	(5)
Constant	0.0871	0.0488	0.1089	0.0776	0.0969
	(0.0289)	(0.1696)	(0.1461)	(0.0301)	(0.0323)
After-tax interest rate	− 0.5769	− 0.5898	− 0.5573	− 0.6006	− 0.5715
	(0.3302)	(0.3553)	(0.3491)	(0.3454)	(0.3598)
Extra regressor		0.0036	− 0.0021	0.2175	− 0.0514
		(0.0158)	(0.0136)	(0.0812)	(0.0288)
		$\ln Y_{\cdot,t}$	$\ln Y_{\cdot,t-1}$	$\Delta \ln Y_{\cdot,t}$	$\Delta \ln Y_{\cdot,t-1}$
J-statistic	19.4284	19.159	18.7774	16.1391	16.0659
	[13, 0.110]	[12, 0.085]	[12, 0.094]	[12, 0.185]	[12, 0.188]

[a]$N = 3249$. Standard errors are in parentheses while the degrees of freedom and p-value for the overidentifying-restrictions test are reported in square brackets. All GMM regressions have a MA(2) weight matrix.

when either $\ln Y_{i,t}$, $\ln Y_{i,t-1}$, or $\Delta \ln Y_{i,t}$ is an extra regressor in Eq. (3). The change in sign on the income variable in columns (2) and (3) compared to column (4) suggests that there might be both a level and a change effect of income on consumption growth. In the results not tabulated, I include both $\ln Y_{i,t}$ and $\Delta \ln Y_{i,t}$ and perform a Wald test of joint significance. The Wald test rejects the null hypothesis that they are jointly zero with a p-value of 0.006. Likewise, composite consumption in Table 5 exhibits excess sensitivity when either $\Delta \ln Y_{i,t}$ or $\Delta \ln Y_{i,t-1}$ is an extra regressor, and when $\ln Y_{i,t}$ and $\Delta \ln Y_{i,t}$ are regressors with a p-value on the Wald test of joint significance equal to 0.007.

The rejection of the hypothesis is robust to both the passbook savings rate and the 3-month T-bill rate as the interest rate $r_{i,t}$. In addition, correlation between errors in the instrument set and the income terms does not seem to be a problem. More specifically, in general, one does not need to worry about measurement error in the explanatory variables provided that it is uncorrelated with any measurement error in the instrument set. However, in the tests reported in Tables 2, 4 and 5 lagged values of disposable income appear as both regressors and instruments such that if measurement error is a problem in the income variables they would not be valid instruments. I tested the validity of the income instruments with the pseudo likelihood ratio test described in Eichenbaum et al. (1988) and could not reject the null that the income variables are valid instruments.[5] Hence, rejection of the permanent-income hypothesis is sensitive to consumption measure.[6]

The estimated interest elasticities of substitution in Tables 4 and 5 require additional explanation as they are either statistically zero or negative. On the surface the negative coefficients are disturbing because in the model of Eq. (3) $1/\alpha_1$ is the coefficient of relative risk aversion and, if negative, implies a implausible nonconcave utility function. However, in the case of predicted consumption, the negative intertemporal substitution parameter in column (1) simply captures the negative influence of current income on consumption growth as shown in column (2). While the negative sign remains in the case of composite consumption, the interest elasticity is, in general, statistically zero.

3.3. Results from the liquidity-constrained tests

Household-based panel data permit the identification of households that might be liquidity constrained. Runkle divides his sample first between

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[5] The validation study of the PSID by Bound et al. (1994) concludes that measurement error in individual earnings data is not significant.

[6] As an additional check that the rejection is not an artifact of the data by the inclusion of oversampled poor households from the SEO subsample, I performed the tests with just the randomly-selected Survey Research Center sample of the PSID. The SRC sample generates qualitatively similar outcomes such that inclusion of the SEO sample does not bias the tests.

homeowners and renters and then as households with more than two months' income in liquid assets versus those households with less. He finds no evidence of liquidity-constrained households in any of the subgroups. I perform a similar division with the current sample and fail to reject the permanent-income hypothesis for homeowners, renters, and households with more than two months' income in liquid assets with all three consumption measures. However, I do find evidence that households with less than two months' income in liquid

Table 6
Tests of the permanent-income hypothesis using food consumption for households with less than two months' income in liquid assets[a]

	(1)	(2)	(3)	(4)	(5)
Constant	0.0575	0.1580	0.1503	0.0617	0.0577
	(0.0132)	(0.0533)	(0.0507)	(0.0141)	(0.0132)
After-tax interest rate	0.1728	0.4373	0.4247	0.2228	0.1585
	(0.1614)	(0.2134)	(0.2111)	(0.1722)	(0.1693)
Extra regressor		− 0.0084	− 0.0077	0.0369	− 0.0062
		(0.0042)	(0.0040)	(0.0435)	(0.0276)
		$\ln Y_{\cdot,t}$	$\ln Y_{\cdot,t-1}$	$\Delta \ln Y_{\cdot,t}$	$\Delta \ln Y_{\cdot,t-1}$
J-statistic	22.5334	19.0868	19.2000	21.7955	22.4980
	[13, 0.048]	[12, 0.086]	[12, 0.084]	[12, 0.040]	[12, 0.032]

[a] $N = 3519$. Standard errors are in parentheses while the degrees of freedom and p-value for the overidentifying-restrictions test are reported in square brackets. All GMM regressions have a MA(1) weight matrix.

Table 7
Tests of the permanent-income hypothesis using predicted consumption for households with less than two months' income in liquid assets[a]

	(1)	(2)	(3)	(4)	(5)
Constant	0.0316	0.2591	0.2482	0.0274	0.0350
	(0.0110)	(0.0593)	(0.0563)	(0.0105)	(0.0116)
After-tax interest rate	− 0.4127	− 0.0027	− 0.0139	− 0.4705	− 0.4021
	(0.1052)	(0.1512)	(0.1496)	(0.0924)	(0.1038)
Extra regressor		− 0.0204	− 0.0195	0.0455	− 0.0260
		(0.0052)	(0.0049)	(0.0500)	(0.0268)
		$\ln Y_{\cdot,t}$	$\ln Y_{\cdot,t-1}$	$\Delta \ln Y_{\cdot,t}$	$\Delta \ln Y_{\cdot,t-1}$
J-statistic	26.2382	11.7935	12.3762	28.1844	24.8265
	[13, 0.016]	[12, 0.462]	[12, 0.416]	[12, 0.005]	[12, 0.016]

[a] $N = 3519$. Standard errors are in parentheses while the degrees of freedom and p-value for the overidentifying-restrictions test are reported in square brackets. All GMM regressions have a MA(1) weight matrix.

Table 8
Tests of the permanent-income hypothesis using composite consumption for households with less than two months' income in liquid assets[a]

	(1)	(2)	(3)	(4)	(5)
Constant	0.0411	0.2541	0.2747	0.0495	0.0645
	(0.0334)	(0.1714)	(0.1479)	(0.0342)	(0.0373)
After-tax interest rate	– 0.8408	– 0.7309	– 0.7344	– 0.8204	– 0.7572
	(0.3713)	(0.3770)	(0.3771)	(0.3850)	(0.3901)
Extra regressor		– 0.0196	– 0.0217	0.1779	– 0.0679
		(0.0157)	(0.0134)	(0.0726)	(0.0273)
		$\ln Y_{\cdot,t}$	$\ln Y_{\cdot,t-1}$	$\Delta \ln Y_{\cdot,t}$	$\Delta \ln Y_{\cdot,t-1}$
J-statistic	19.3900	17.3528	16.2262	12.5810	12.1599
	[13, 0.111]	[12, 0.137]	[12, 0.181]	[12, 0.400]	[12, 0.433]

[a] $N = 2685$. Standard errors are in parentheses while the degrees of freedom and p-value for the overidentifying-restrictions test are reported in square brackets. All GMM regressions have a MA(2) weight matrix.

assets are liquidity constrained using either the predicted or composite consumption measures.

The results, reported in Tables 6–8, follow a similar pattern as the results from Tables 2, 4 and 5. There is limited evidence of liquidity-constrained behavior with food consumption using $\ln Y_{i,t}$ or $\ln Y_{i,t-1}$. However, the evidence is not robust because, unlike the tests with predicted and composite consumption, the Wald tests that $\ln Y_{i,t}$ and $\Delta \ln Y_{i,t}$ or $\ln Y_{i,t-1}$ and $\Delta \ln Y_{i,t}$ are jointly insignificant are not rejected with p-values of 0.12. Consequently, testing for liquidity-constrained households using the less-comprehensive food consumption measure leads to the false acceptance of the permanent-income hypothesis for low liquid-asset income households.

4. Conclusions

I propose a composite consumption measure for PSID data that is based on household wealth. While measurement error is a more significant problem with the composite consumption measure compared to food consumption or Skinner's (1987) predicted consumption, Skinner's measure may be unstable if the relative prices of goods change, and many of the components needed to construct his measure are no longer collected by the PSID. I further show that tests of the permanent-income hypothesis and for liquidity-constrained households are sensitive to the choice of consumption measure in the PSID. In particular, food consumption fails to reject the hypothesis, but it is rejected with broad consumption measures such as predicted consumption and composite consumption.

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