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Habit persistence and durability in aggregate consumption

Empirical tests

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Habit persistence in preferences and durability of consumption goods both imply time-nonseparability in the derived utility for consumption expenditures. We study a simple model with both effects. Lagged consumption expenditures enter the Euler equation, where habit persistence implies that their coefficients are negative and durability implies positive coefficients. Estimating the sign of the coefficients addresses the question of which effect is dominant. Earlier empirical work on monthly data supports the durability of consumption expenditures. We find evidence in monthly, quarterly, and annual data that habit persistence dominates the effect of durability.

1. Introduction

The consumption-based asset pricing model tested by Hansen and Singleton (1982, 1983), Ferson (1983), Grossman, Melino, and Shiller (1987), Breeden, Gibbons, and Litzenberger (1989), and others assumes that the utility function is time- and state-separable and that the consumption good is

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nondurable. Habit persistence in consumption preferences and durability of consumption goods both imply time-nonseparability in the derived utility for consumption expenditures.

We illustrate the combined effects of durability in consumption expenditures and of habit persistence in preferences, using a simple theoretical model in which the durable good depreciates by a fixed proportion in each period. Habit persistence is modeled by assuming that a consumer's subsistence level is a weighted sum of the past flows of consumption services. In this model, both current and lagged consumption expenditures enter the Euler equation. Habit persistence implies that the coefficients on the lagged expenditures are negative, whereas durability implies positive coefficients. If both effects are present, the sign of the coefficients indicates which is dominant.

In earlier work, Dunn and Singleton (1986), Eichenbaum, Hansen, and Singleton (1988), and Eichenbaum and Hansen (1990) estimated a positive coefficient on lagged consumption in Euler equations for monthly data. They interpret their results as evidence of durability in consumption expenditures. These studies used lagged consumption and returns as the predetermined information variables in the model. Monthly consumption data are suspect, however, for at least two reasons. First, measurement error, which can induce negative autocorrelation in consumption growth, is likely to be proportionately more important in monthly data. Second, components of monthly consumption expenditures are calculated by interpolation, which may induce positive autocorrelation. Durability of consumption expenditures induces negative autocorrelation. For example, a consumer purchasing an automobile in one period is unlikely to purchase another for several periods. Habit persistence induces positive autocorrelation in consumption growth, since the consumer maximizes utility by smoothing consumption more than would be optimal with time-separable preferences. Therefore, spurious positive or negative autocorrelation may lead to the erroneous conclusion that habit persistence dominates or is dominated by durability of consumption expenditures.

We find that evidence of the importance of durability in monthly consumption data is weak. When we experiment with different predetermined information variables or allow for autocorrelation in the error terms, as would be implied by time-aggregation of the consumption data, the evidence suggests that habit persistence is dominant.

When we extend the investigation to quarterly and annual data, we find that habit persistence dominates durability at these frequencies as well. The evidence of habit persistence in the quarterly and annual data is robust to the autocorrelation assumption and the choice of instrumental variables. Durability could have a sufficiently short half-life in relation to habit persistence that it is suppressed in the quarterly and annual data. In a calibration exercise to explain Mehra and Prescott's (1985) equity premium puzzle, Constantinides (1990) finds evidence that habit persists longer than a year.

Our results are reinforced by the recent work of Hansen and Jagannathan (1991) and Gallant, Hansen, and Tauchen (1990), who study the moment inequality restrictions implied by the Euler equations. Winder and Palm (1989) estimate a linearized form of the Euler equation and find support for habit persistence in the Netherlands. Heaton (1990) considers an explicit consumption process, assumes that the interest rate is constant, and formally models the time-averaging of monthly and quarterly consumption data in a linearized version of the Euler equation. He finds evidence for short-lived durability in consumption expenditures and some evidence for habit persistence. Backus, Gregory, and Telmer (1990) find that habit persistence helps to account for the variability of expected returns on currencies.

The paper is organized as follows. The model is stated and the Euler equation that incorporates nonseparable preferences and durability of consumption expenditures is derived in section 2. Our method is described in section 3. In section 4 we discuss the monthly, quarterly, and annual consumption expenditures data, the asset-returns data, and the predetermined instrumental variables. The main empirical results, presented in tables 4–7, are discussed in section 5. The robustness of the empirical results is examined further in section 6, and in section 7 we reconsider the equity premium puzzle. In the concluding section 8, we offer suggestions for future work. An appendix illustrates the interpretation of the concavity parameter of the representative agent's utility function under time-nonseparability. We argue that the parameter more closely approximates the risk-aversion coefficient than the inverse of the elasticity of consumption with the interest rate.

2. The model

We consider a single-good economy in discrete time. Expenditures on the good at time t by a representative consumer are denoted by c_t. The good is durable, and durability is modeled as in Dunn and Singleton (1986), Eichenbaum, Hansen, and Singleton (1988), and Eichenbaum and Hansen (1990). Each period the new expenditures c_t produce a flow of consumption services $\delta_\tau c_t$ in period $t+\tau$, $\tau \geq 0$, where $\delta_\tau \geq 0$ and $\sum_{\tau=0}^{\infty}\delta_\tau = 1$. The total flow of consumption services at time t is given by

$$c_t^F = \sum_{\tau=0}^{\infty} \delta_\tau c_{t-\tau}. \tag{1}$$

The representative consumer's utility is defined over the flow of services c_t^F.

We model habit persistence with a time-nonseparable von Neumann–Morgenstern utility function

$$(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t\left(c_t^F - h\sum_{s=1}^{\infty}a_s c_{t-s}^F\right)^{1-A}, \tag{2}$$

where $A > 0$, $a_s \geq 0$, $\Sigma_{s=1}^{\infty}a_s = 1$, and $h \geq 0$. The time-preference parameter is β. The habit parameter h represents the fraction of the weighted sum of lagged consumption flows that establishes a subsistence level of consumption. If $h = 0$, the utility function is time-separable in consumption flows (but not in consumption expenditures, unless $\delta_\tau = 0$, $\tau \geq 1$).

Ryder and Heal (1973) study the optimal consumption policy when the utility at time t is defined as a concave function of the consumption flow at time t and a weighted sum of the lagged consumption flows. Sundaresan (1989), Constantinides (1990), and Novales (1990) study special cases in which the utility at time t is a power of the difference between the consumption flow at time t and a fraction of a weighted sum of lagged consumption flows, as in (2).

In the case of time-separable preferences over consumption services ($h = 0$), the concavity parameter A is the relative-risk-aversion coefficient (RRA) and the inverse of the elasticity of consumption with respect to the interest rate. With habit persistence ($h > 0$), we show in the appendix that the parameter A approximately equals the RRA coefficient but may differ substantially from the inverse of the elasticity of consumption.

We combine (1) and (2) and write the utility function as

$$(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}, \tag{3}$$

where

$$\begin{aligned} C_t &\equiv \sum_{\tau=0}^{\infty}\delta_\tau c_{t-\tau} - h\sum_{s=1}^{\infty}\sum_{\tau=0}^{\infty}a_s\delta_\tau c_{t-\tau-s} \\ &= \delta_0\sum_{\tau=0}^{\infty}b_\tau c_{t-\tau} \end{aligned} \tag{4}$$

and

$$b_0 = 1,$$

$$b_\tau = \left(\delta_\tau - h\sum_{i=1}^{\tau}a_i\delta_{\tau-i}\right)/\delta_0, \qquad \tau \geq 1.$$

It is instructive to consider an example with $\delta_\tau = (1-\delta)\delta^\tau$, $0 \le \delta < 1$, and $a_s = (1-\alpha)\alpha^{s-1}$, $0 \le \alpha < 1$. Then the coefficients on the lagged expenditures become

$$b_\tau = \left[1 - \frac{(1-\alpha)h}{\delta-\alpha}\right]\delta^\tau + \frac{(1-\alpha)h\alpha^\tau}{\delta-\alpha}, \qquad \tau \ge 1. \tag{5}$$

If expenditures are not durable ($\delta = 0$), we obtain $b_\tau = -(1-\alpha)h\alpha^{\tau-1}$ and the coefficients b_τ are negative for $\tau \ge 1$. In the absence of habit persistence ($h = 0$) but with durability, we obtain $b_\tau = \delta^\tau$ and the coefficients are positive. When both habit persistence and durability are present, the coefficients b_τ are positive or negative depending on the relative magnitudes of the durability parameter δ and the habit persistence parameters h and α. If $\delta \ge \alpha + h(1-\alpha)$, the coefficient b_τ is positive for all τ; if $\delta \le h(1-\alpha)$, then b_τ is negative for all $\tau \ge 1$; finally, if $h(1-\alpha) < \delta < \alpha + h(1-\alpha)$, b_τ is positive for recent lags and negative for distant ones. The example illustrates the opposing effects of habit persistence and durability on the coefficients of lagged consumption expenditures. In particular, assuming exponential decay for both habit and durability, it shows that if habit persistence dominates durability ($b_\tau < 0$) at a given lag τ, then habit must dominate durability at all greater lags ($b_j < 0$ for all $j > \tau$).

To derive the stochastic Euler equation in the general case, we consider a reduction of the representative consumer's expenditures in period t from c_t to $c_t - \varepsilon$, where $|\varepsilon|$ is small. The investment of ε in an asset with (stochastic) return R_{t+1} over one period increases the consumption expenditures in period $t+1$ from c_{t+1} to $c_{t+1} + \varepsilon R_{t+1}$. The rational consumer takes into account the effect of the changes in consumption expenditures in periods t and $t+1$ on the flow of consumption services and on the subsistence level in all future periods through (4) and calculates the change in $C_{t-\tau}$, C_t, and $C_{t+\tau}$, $\tau \ge 1$, as

$$\begin{aligned}
&\partial C_{t-\tau}/\partial\varepsilon = 0, \qquad \tau \ge 1, \\
&\partial C_t/\partial\varepsilon = -\delta_0, \\
&\partial C_{t+\tau}/\partial\varepsilon = (b_{\tau-1}R_{t+1} - b_\tau)\delta_0, \qquad \tau \ge 1.
\end{aligned} \tag{6}$$

Optimality requires that the expectation in period t of the utility of the consumption flows is maximized at $\varepsilon = 0$, that is,

$$\partial/\partial\varepsilon|_{\varepsilon=0}\left[(1-A)^{-1}\sum_{\tau=-t}^{0}\beta^{t+\tau}C_{t+\tau}^{1-A} + \mathrm{E}_t\left[(1-A)^{-1}\sum_{\tau=1}^{\infty}\beta^{t+\tau}C_{t+\tau}^{1-A}\right]\right]$$
$$= 0. \tag{7}$$

We combine (6) and (7), simplify, and obtain the Euler equation

$$\mathrm{E}_t\left[\sum_{\tau=1}^{\infty}\beta^{\tau}(C_{t+\tau}/C_t)^{-A}(b_{\tau-1}R_{t+1}-b_{\tau})-1\right]=0, \tag{8}$$

where C_t is defined by (4).

In the absence of habit persistence ($h=0$) and durability ($\delta_\tau=0,\ \tau\geq 1$), we obtain $b_\tau=0,\ \tau\geq 1$, and the Euler equation (8) becomes the time- and state-separable model examined by Hansen and Singleton (1982):

$$\mathrm{E}_t\left[\beta(c_{t+1}/c_t)^{-A}R_{t+1}-1\right]=0. \tag{9}$$

We consider a sequence of nested models of the Euler equation, starting with the time-separable model ($b_\tau=0,\ \tau\geq 1$) and proceeding to a model with a one-lag specification ($b_\tau=0,\ \tau\geq 2$) and a model with a two-lag specification ($b_\tau=0,\ \tau\geq 3$). Formally, the one-lag model captures habit persistence based on one lag in consumption, in the absence of durability; or it captures durability of one period only, in the absence of habit persistence. The two-lag model captures habit persistence based on two lags in consumption, in the absence of durability; or durability of two periods, in the absence of habit persistence; or habit persistence based on one lag in consumption and durability of one period.

3. Empirical method

We test the Euler equation and estimate the model parameters using Hansen's (1982) generalized method of moments (GMM). Eq. (8) defines an error term u_{it} for each asset i, $i=1,\dots,N$, such that $\mathrm{E}_t[u_{i(t+1)}]=0$, where $\mathrm{E}_t[\cdot]$ denotes the conditional expectation given information at time t. With a set of L instruments, z_{jt}, $j=1,\dots,L$, known to the market at time t, we obtain $\mathrm{E}[u_{t+1}\mid z_t]=0$ and therefore $\mathrm{E}[u_{t+1}\ z_t']=0$, where u_{t+1} is the vector of N error terms and z_t is the vector of L instruments. Given N assets and L instruments, there are NL orthogonality conditions. The GMM estimates are based on minimizing the quadratic form $g'Wg$ where g is the NL vector of the elements of $(1/T)\Sigma_t u_{t+1}z_t'$ and W is the inverse of a consistent estimate of the covariance matrix of these orthogonality conditions. Hansen (1982) discusses the weighting matrix W and provides conditions under which the parameter estimates are consistent and asymptotically normal and the minimized value of the quadratic form is asymptotically chi-square under the null hypothesis. The model is overidentified provided that the number of orthogonality conditions, NL, exceeds the number of parameters. The minimized quadratic form provides a test statistic for the goodness-of-fit of the model; the number of degrees of freedom is the number of orthogonality conditions minus the number of parameters. The parameters are $\beta, A, \{b_\tau\}$.

In the two-lag model if we choose $A = 0$, the Euler equation (8) implies $u_{t+1} = (1 + b_1\beta + b_2\beta^2) - \beta R_{t+1}(1 + b_1\beta + b_2\beta^2)$. If we also choose the parameters such that $1 + b_1\beta + b_2\beta^2 = 0$, we obtain a trivial solution to the Euler equation. Following Eichenbaum and Hansen (1990), we note that the orthogonality condition $\mathrm{E}\{u_{t+1} \mid Z_t\} = 0$ holds if u_{t+1} is divided by a constant. We divide the Euler equation by the term $(1 + b_1\beta + b_2\beta^2)$ in order to avoid trivial solutions.

In the time-separable model, u_t is a function of the variables R_t, c_{t-1}, and c_t, which are known at time t. The Euler equation therefore implies that $\mathrm{E}[u_{t+s} \mid u_t] = 0$, $s \geq 1$, and we say that u_t follows an MA(0) process. The time-separable model implies the null hypothesis,

$$\mathrm{H}_0\colon \quad b_s = 0, \quad s \geq 1, \quad \text{with an MA(0) error term } u_t.$$

In the one-lag model ($b_s = 0$, $s \geq 2$), u_t is a function of R_t, c_{t-2}, c_{t-1}, c_t, and c_{t+1}. Since c_{t+1} is not in the time t information set, the model does not imply that $\mathrm{E}[u_{t+1} \mid u_t] = 0$, but it does imply that $\mathrm{E}[u_{t+s} \mid u_t] = 0$, $s \geq 2$. We say that u_t in this case follows an MA(1) process. In general, the model implies that the error term u_t will be MA(q), where q is the smallest number such that $b_j = 0$ for all $j > q$. The one-lag model therefore implies the hypothesis,

$$\mathrm{H}_a\colon \quad b_s = 0, \quad s \geq 2, \quad \text{with MA(1) error term } u_t.$$

The autocorrelation of the error term under H_a is related to the parameter b_1 in a complex way. The weighting matrix is adjusted to account for the moving-average terms, as described by Hansen (1982). When b_1 is zero, the model implies that the autocorrelation of the error becomes zero, hence the null hypothesis H_0.

We model the consumer's decisions at fixed intervals, and measure asset returns and consumption over the same intervals. Consumption decisions may actually occur more often. If the decisions are made within the observation interval and the measured consumption expenditures are the sum of the expenditures over the interval, the consumption data are said to be time-aggregated. Formally modeling time aggregation in the Euler equation is difficult and theoretical results for time aggregation are only available in the literature, imposing a first-order approximation on the marginal utility function. Heaton (1990) studies monthly data using a first-order approximation of the Euler equation in this paper. He uses lagged consumption and dividends as the information set and models time aggregation. His results are consistent with our estimates of b_1, and they suggest that time aggregation may not be a very important influence on the estimates. We conduct several experiments, however, to assess its likely importance further.

Using a linear approximation to the Euler equation, we can show that one effect of time aggregation is to increase by one the order of the moving-average process followed by u_t. In the general, nonlinear Euler equation, the results can be more complex. Therefore, under time aggregation, the residuals may appear to behave like a higher-order MA process even if the nonseparability parameter b_1 is zero. Time aggregation can also induce a spurious correlation between the error terms and the information set for time t. Therefore, variables in the market's information set at t that were not present at $t-1$ may not be valid instruments for the equation $\mathrm{E}_t[u_{t+1}] = 0$.

Time aggregation is not the only feature of the aggregate consumption data that creates methodological problems. Other features could induce autocorrelation in the error terms and spurious correlation between the error terms and the instruments. Examples include imperfect timing and interpolation of the consumption data, measurement errors, and seasonal adjustment. A rejection of H_0 could be interpreted as evidence that these other features are important, rather than as evidence that b_1 is not zero. We therefore examine the modified null hypothesis,

$$\mathrm{H}_0'\colon \quad b_1 = 0, \quad \text{with an MA(1) error term } u_t,$$

against the alternative hypothesis H_a.

We further assess the sensitivity of our results to these data issues by conducting experiments in which the order of the MA process of the errors is varied and in which the instruments for the information set at time t either admit or do not admit the most recent lagged values of the variables.

The parameter estimates and statistical tests using the GMM are justified from asymptotic distribution theory. There is a natural concern about the properties of these procedures in small samples. Tauchen (1986) and Kocherlakota (1990) provide simulation evidence for the time-separable model. Tauchen finds that the test statistics perform well with as few as 50 annual observations, although there is a slight tendency to reject the model too infrequently. Kocherlakota (1990), using a different set of parameter values, finds cases where the model is rejected too often. When the tests exploit only unconditional moment restrictions, that is, when the instrument is a constant vector of ones (a case we examine below), the test statistics perform well in small samples. Both of these studies find that the coefficient estimates and their standard errors can be unreliable in small samples.

Although we report the coefficient estimates and their asymptotic standard errors, we stress that their reliability cannot be assessed until simulation studies of the finite-sample properties of the GMM become available for nonseparable consumption models. We therefore refrain from deriving detailed implications from the models that depend on the point estimates of the coefficients.

Ferson and Foerster (1991) study the finite-sample properties of the GMM in a different asset pricing context that involved regressions with nonlinear, cross-equation restrictions. They find that a two-stage GMM approach, as described in Hansen and Singleton (1982), tends to reject the model too often in larger systems, and an iterated GMM approach provides more accurate test statistics. We use an iterated GMM approach in this study.[1]

4. The data

We study the returns on Treasury bills, bonds, and value-weighted portfolios of common stocks traded on the New York Stock Exchange (NYSE). The stocks are grouped into size deciles, based on the market value of equity outstanding at the beginning of each year. We examine a subset of the deciles: deciles 1, 5, and 10. Decile 1 represents the common stocks of small firms and decile 10 represents large firms. Three portfolios are chosen to keep the number small, while capturing most of the stock-return behavior that would be reflected in a design using all ten deciles or value- and equally-weighted stock indices. We include a long-term government bond return and a strategy of rolling over one-month Treasury bills. Thus, a total of five portfolio returns are examined. All of the asset-return data are from the Center for Research in Security Prices of the University of Chicago (CRSP).

Our primary measure of consumption is real, per-capita expenditures on consumer nondurable goods. These data are seasonally adjusted by the Commerce Department, using the X-11 seasonal adjustment program. The quarterly data are in real terms, as reported by Data Resources Inc. (DRI). The annual consumption data are from the Commerce Department's *Business Statistics*, 1959 edition, and DRI. The annual data for 1929–1949 are spliced into the annual sums of the quarterly Commerce Department data, in real, per-capita terms, using 1949 levels as the splicing factor. Monthly data are obtained from Citibase. The real expenditure totals are divided by the population to obtain the per-capita, real consumption series. The population figure is the total United States residential population (excluding armed forces abroad) from *Statistical Abstract of the U.S.*, DRI, and Citibase.

We examine data for durable goods expenditures and for seasonally unadjusted consumption in some experiments. The seasonally unadjusted data are the quarterly nominal expenditures from DRI, divided by the

[1]Specifically, we construct the weighting matrix W using the parameter estimates from the nth stage, use this matrix to find parameters for stage $n+1$ that minimize the quadratic form, and then use the new parameters to update the weighting matrix. The iterations continue until a minimum value of the quadratic form is obtained.

population and deflated by a seasonally unadjusted component of the Consumer Price Index (CPI). We use the CPI components because price deflators for personal consumption expenditures are available to us only in seasonally adjusted form. We use the seasonally unadjusted components of the CPI for consumer nondurable goods to deflate the nominal nondurables expenditure totals, and the CPI for consumer durables to deflate the durable goods expenditures. (In other experiments, we used the overall CPI as the deflator for both categories, and the results were similar.)

The real asset returns are the nominal returns deflated by the price deflator for the measure of consumption in a given model. For example, when we use nondurable goods expenditure data we deflate the asset returns by the deflator for consumer nondurable goods. When we use the seasonally unadjusted consumer durables expenditure data, we deflate the returns by the seasonally unadjusted Consumer Price Index for consumer durables.

Previous studies of consumption-based models use lagged values of consumption and returns to estimate the parameters and test the Euler equations. Measurement errors and other data problems can result, however, in spurious correlation between the consumption and real returns and their lagged values that leads to spurious rejections of the Euler equations and biases the parameter estimates. We use lagged consumption and returns in some experiments for comparison and to check the sensitivity of our results. We focus mainly on instruments that are distinct from the lagged values, but can predict both asset returns and measures of consumption growth. Such instruments should provide powerful tests of the Euler equation restrictions. The instrumental variables are summarized in table 1.

VWYLD(-1) is the average dividend yield on a value-weighted index of common stocks traded on the NYSE, provided by CRSP. The dividend yield is the sum of the most recent year's dividends divided by the price level on the last trading day of the period (month, year, or quarter). The symbol (-1) indicates that a variable is lagged one period in relation to the date of the asset-return realization in the Euler equation. For example, the dividend yield used to predict returns for the first quarter uses the price level at the end of the previous December and dividends over the previous year. Using the annual dividends avoids the seasonality of dividend payments. Dividend yields are a component of stock returns, so the *ex ante* dividend yield is a natural instrument for capturing variation in expected stock returns. Campbell and Shiller (1988), Fama and French (1988), and others find that dividend yields predict future stock and bond returns.

DIVDIFF(-1) is the dividend yield of the CRSP equally-weighted stock index minus the yield of the value-weighted index. Movements in this variable over time reflect the differences between the dividend yields of small and large firms. We find that the difference contributes additional explanatory power, given the level of the yield, in regressions for the future-returns and consumption measures.

Table 1

Summary of the financial instrumental variables used to estimate and test models with habit persistence and durability in aggregate consumption.

Symbol[a]	Definition	Source
VWYLD(−1)	Dividend yield on the CRSP value-weighted stock index.	CRSP[b]
DIVDIFF(−1)	Dividend yield on the CRSP equally-weighted stock index, minus the dividend yield of the value-weighted stock index.	CRSP
TB1MO	Nominal one-month Treasury bill rate.	CRSP
PI(−1)	Inverse of the price level index for the smallest decile of common stocks, multiplied by the average of the price level over the previous year.	CRSP
GIPX(−1)	Annual growth rate of the U.S. industrial production index.	FR[c] Bulletin
LSLOPE(−1)	Average monthly yield-to-maturity of corporate bonds rated Aaa by Moody's Investor Services, minus the one-month Treasury bill rate.	FR Bulletin
SSLOPE(−1)	Three-month Treasury bill rate less the one-month Treasury bill rate.	CRSP
CBPREM(−1)	Average monthly yield-to-maturity of corporate bonds rated Baa by Moody's Investor Services, minus the Aaa corporate bond yield.	FR Bulletin

[a]The (−1) symbol indicates that the variable is lagged one period in relation to the asset returns in the Euler equation.
[b]Center for Research in Security Prices at the University of Chicago.
[c]Federal Reserve.

TB1MO is the nominal one-month Treasury bill rate. The ability of short-term bills to predict monthly returns of bonds and stocks is documented by Fama and Schwert (1977) and others. The nominal bill rate is from the CRSP RISKFREE files. The one-month bill is for the first month of the period over which the return to be predicted is measured. For example, to predict the first quarter returns, we use the bill quoted on the last trading day of December that is the closest to one month to maturity.

PI(−1) is a measure of the detrended price level for the smallest decile of common stocks. This is the inverse of the price index level, relative to the average level over the preceding twelve months. Keim and Stambaugh (1986) study a similar variable and find that it predicts both bond and stock returns. This variable may capture the reversion of expected returns to their long-run means. Mean reversion implies that if stock returns are below average (so that prices are relatively low), conditional expected returns are higher than average.

GIPX(− 1) is the continuously compounded annual growth rate of an index of U.S. industrial production.

SSLOPE(− 1) is the three-month Treasury bill rate, minus the one-month return of a one-month bill. This is one of three instruments that decompose risky debt yields into short-term and long-term term-structure slope variables, and an *ex ante* default premium for corporate debt. Fama (1984), Campbell (1987), and Stambaugh (1988) find that short-term measures of the term structure can predict bond returns of different maturities and stock returns.

LSLOPE(− 1) is the long-term slope, measured as the lagged value of the yield-to-maturity of Aaa corporate bonds, minus the one-month Treasury bill rate.

CBPREM(− 1) is the lagged value of the average monthly yield-to-maturity of corporate bonds rated Baa by Moody's Investor Services, minus the lagged value of the Aaa corporate bond yield. Keim and Stambaugh (1986) find that a yield spread has predictive power for bond and stock returns.

The predetermined variables follow empirical studies that document their ability to predict the returns of portfolios similar to the ones we study. Statistical inference is complicated if the variables are the result of collective 'data snooping' by a series of researchers. We do not attempt to account formally for these effects in our analysis. If there is spurious predictability, however, it should be difficult for the Euler equation to 'explain' it. Such a bias is conservative, given our result that habit persistence helps to explain the predictability through the Euler equation. [See Lo and MacKinlay (1990) for an analysis of data snooping in financial models.]

Tables 2 and 3 present summary statistics for the basic data. The sample period of the quarterly analysis is 1948–1986; the annual data are for 1929–1986; monthly data cover 1959–1986. For annual data we use a subset of the financial instrumental variables. These are *VWYLD*, *DIVDIFF*, *TB1MO*, *P1*, and *SSLOPE*. Table 2 shows that some of the autocorrelations of the instruments are above 0.9 (the quarterly *CBPREM*, *VWYLD*, *DIVDIFF*, and *TB1MO*, and also *GIPX* in monthly data). For *VWYLD*, *P1*, and *DIVDIFF*, this is expected, given the overlapping nature of the numerators. *GIPX* is an annual growth rate, and overlaps in the monthly and quarterly data. The autocorrelations decay toward zero at longer lags for all the variables. Table 3 shows the contemporaneous correlations among the instruments, which suggest that none of the instruments are redundant. Only three of the correlations among the quarterly and two among the annual instruments exceed 0.5, and the largest is 0.67 (0.69). In monthly data, five of the correlations exceed 0.5 and the largest is 0.74.

We examine time-series regressions, using the instruments to predict the future returns of the common stock and bond portfolios and the future growth rates of consumption. The asset returns are measured in excess of the

Table 2

Summary statistics for all variables for quarterly, annual, and monthly observations.

Consumption growth rates for real per-capita expenditures are measured in decimal fraction units per period. Returns are arithmetic nominal rates of return, in decimal fraction units per period. The one-month bill return corresponds to a strategy of rolling over one-month Treasury bills each month. Dec*N* refers to a value-weighted portfolio of common stocks of the *N*th size decile. Dec1 denotes the decile of the smallest firms and Dec10 the largest firms. *CBPREM* is the annualized yield-to-maturity of corporate bonds rated Baa by Moody's Investor Services, minus the yield of Aaa-rated bonds. *TB1MO* is the continuously compounded rate of return of a one-month Treasury bill. *VWYLD* is the dividend yield on the CRSP value-weighted index, measured as the previous twelve months' dividend payments divided by the level of the index. *DIVDIFF* is the difference between the dividend yield of the CRSP equally-weighted index and that of the value-weighted index. *LSLOPE* is the Aaa corporate bond yield, minus the one-month Treasury bill rate. *SSLOPE* is the difference between the three-month and the one-month Treasury bill yield. *GIPX* is the continuously compounded annual growth rate of the index of industrial production. All instruments are measured in decimal fraction units except *PI*. *PI* is the inverse of the index level of prices of the smallest decile of common stocks on the New York Stock Exchange, relative to the average level for the preceding twelve months. X-11 indicates data seasonally adjusted by the Commerce Department using the X-11 program. DSA indicates consumption data seasonally adjusted by the authors using dummy variables. NSA denotes not seasonally adjusted.

Variable	Mean	Std. dev.	Autocorrelations ρ_1	ρ_2	ρ_3	ρ_4	ρ_8	ρ_{12}	ρ_{24}	ρ_{36}
			Part I: Quarterly data							
			Real consumption growth rates (1948.2Q–1986.2Q, 153 observations)							
Nondurables (X-11)	0.003	0.009	0.07	0.08	0.12	−0.05	−0.17	−0.01	−0.01	−0.16
Durables (X-11)	0.009	0.042	−0.09	0.15	−0.14	−0.03	−0.20	−0.09	0.03	0.02
Nondurables (DSA)	0.001	0.022	−0.31	−0.21	−0.14	0.56	0.40	0.50	0.31	0.11
Durables (DSA)	0.002	0.052	−0.24	0.22	−0.27	0.19	−0.02	0.01	−0.01	0.01
			Asset returns (1947.1Q–1986.4Q, 160 observations)							
1-month bill	0.012	0.008	0.96	0.92	0.90	0.87	0.75	0.68	0.51	0.38
Government bond	0.012	0.049	−0.06	0.07	0.15	0.06	−0.06	−0.01	0.00	−0.11
Stocks: Dec1	0.047	0.133	0.01	−0.11	−0.08	0.15	−0.04	0.13	−0.07	0.15
Stocks: Dec5	0.037	0.099	0.10	−0.10	−0.06	−0.01	−0.07	0.03	−0.17	0.09
Stocks: Dec10	0.029	0.071	0.13	−0.06	−0.05	−0.01	−0.07	0.05	−0.13	0.09
			Instrumental variables							
CBPREM	0.009	0.005	0.92	0.85	0.80	0.73	0.54	0.41	0.39	0.26
TB1MO	0.047	0.031	0.91	0.90	0.88	0.85	0.75	0.67	0.50	0.37
VWYLD	0.041	0.011	0.94	0.88	0.83	0.78	0.64	0.55	0.08	−0.19

Table 2 (continued)

Variable	Mean	Std. dev.	Autocorrelations ρ_1	ρ_2	ρ_3	ρ_4	ρ_8	ρ_{12}	ρ_{24}	ρ_{36}
			Instrumental variables (continued)							
DIVDIFF	−0.002	0.005	0.92	0.87	0.79	0.74	0.53	0.44	0.24	0.14
LSLOPE	0.019	0.013	0.66	0.61	0.55	0.45	0.17	−0.04	0.02	0.02
GIPX	0.037	0.066	0.81	0.47	0.10	−0.25	−0.16	−0.05	−0.00	−0.00
PI	0.922	0.149	0.60	0.27	0.06	−0.02	−0.17	0.08	−0.29	0.00
SSLOPE	0.003	0.005	0.15	0.35	0.21	0.50	0.37	0.07	0.15	0.06
			Inflation rates (1948.2Q–1986.2Q, 153 observations)							
Nondurables (X-11)	0.009	0.014	0.69	0.54	0.48	0.31	0.09	0.18	0.23	−0.01
Durables (X-11)	0.008	0.012	0.08	0.09	0.14	0.08	0.36	0.00	0.25	0.03
Nondurables (NSA)	0.009	0.013	0.49	0.36	0.38	0.38	0.08	0.16	0.20	−0.03
Durables (NSA)	0.008	0.013	0.37	0.37	0.26	0.46	0.43	0.31	0.25	0.13
			Part II: Annual data							
			Real consumption growth rates (1930–1986, 57 observations)							
Nondurables	0.013	0.031	0.41	0.11	−0.09	−0.21	−0.03	−0.14	−0.06	−0.09
Durables	0.028	0.150	0.28	−0.07	−0.31	−0.23	−0.03	−0.00	−0.06	−0.26
			Asset returns (1926–1986, 61 observations)							
1-month bill	0.035	0.034	0.92	0.85	0.80	0.75	0.70	0.64	0.59	0.53
Government bond	0.046	0.085	0.10	0.02	0.11	0.10	−0.14	0.09	−0.10	−0.00
Stocks: Dec1	0.228	0.469	0.15	−0.15	−0.13	−0.34	−0.14	−0.14	−0.03	0.07
Stocks: Dec5	0.153	0.290	0.04	−0.15	−0.09	−0.20	−0.03	−0.14	0.01	−0.04
Stocks: Dec10	0.108	0.195	0.04	−0.21	−0.04	−0.14	−0.01	−0.07	0.10	0.05
			Instrumental variables (1926–1986, 61 observations)							
TBIMO	0.033	0.029	0.91	0.84	0.79	0.76	0.71	0.66	0.59	0.53
VWYLD	0.045	0.014	0.62	0.30	0.20	0.21	0.29	0.30	0.22	0.15
DIVDIFF	−0.004	0.008	0.70	0.39	0.28	0.32	0.35	0.33	0.30	0.19
PI	1.004	0.334	0.26	0.02	−0.10	−0.21	−0.16	0.16	0.11	0.09
SSLOPE	0.003	0.006	0.59	0.38	0.12	0.20	0.34	0.32	0.22	0.13

			Inflation rates (1930–1986, 57 observations)							
Nondurables	0.041	0.071	0.66	0.22	−0.05	−0.12	−0.04	0.13	0.12	0.03
Durables	0.035	0.149	0.43	0.07	−0.24	−0.27	0.01	0.08	−0.04	−0.22
			Part III: Monthly data							
			Real consumption growth rates (February 1959–December 1986, 335 observations)							
Nondurables (X-11)	0.001	0.008	−0.36	0.03	0.11	−0.08	0.01	−0.04	−0.24	−0.06
Services (X-11)	0.002	0.004	−0.16	0.00	−0.03	−0.08	0.10	−0.04	−0.05	0.09
Durables (X-11)	0.004	0.030	−0.11	−0.09	−0.17	−0.01	0.05	0.11	−0.12	−0.02
Total (X-11)	0.002	0.006	−0.14	−0.01	−0.01	0.00	0.05	0.02	−0.21	−0.06
			Asset returns (January 1959–December 1986, 336 observations)							
1-month bill	0.005	0.002	0.96	0.93	0.90	0.88	0.83	0.76	0.58	0.45
Government bond	0.005	0.029	0.02	0.02	−0.15	0.06	0.01	−0.05	−0.07	0.01
Stocks: Dec1	0.015	0.071	0.08	−0.02	−0.03	−0.01	−0.12	0.23	0.03	0.10
Stocks: Dec5	0.012	0.054	0.12	−0.03	0.02	0.04	−0.13	0.10	0.01	0.00
Stocks: Dec10	0.008	0.041	−0.03	−0.02	0.05	0.08	−0.02	0.07	−0.02	−0.03
			Instrumental variables (December 1958–November 1986, 336 observations)							
CBPREM	0.011	0.005	0.97	0.93	0.91	0.89	0.78	0.67	0.47	0.31
TB1MO	0.059	0.029	0.96	0.93	0.90	0.88	0.84	0.77	0.58	0.45
VWYLD	0.038	0.009	0.98	0.96	0.94	0.92	0.82	0.75	0.67	0.66
DIVDIFF	−0.003	0.004	0.97	0.96	0.94	0.92	0.85	0.81	0.54	0.34
LSLOPE	0.020	0.015	0.89	0.80	0.73	0.67	0.57	0.40	0.08	−0.16
GIPX	0.039	0.058	0.96	0.90	0.81	0.72	0.37	−0.02	−0.33	−0.07
PI	2.548	47.67	0.07	0.11	0.36	−0.11	−0.03	−0.05	0.03	0.01
SSLOPE	0.004	0.004	0.34	0.30	0.20	0.19	0.22	0.43	0.26	0.02
			Inflation rates (February 1959–December 1986, 335 observations)							
Nondurables (X-11)	0.004	0.005	0.36	0.36	0.23	0.31	0.27	0.14	0.07	−0.04
Durables (X-11)	0.003	0.005	0.01	0.23	0.11	0.13	0.13	0.11	−0.02	0.03
Services (X-11)	0.004	0.003	0.58	0.65	0.64	0.59	0.60	0.60	0.51	0.45
Total (X-11)	0.004	0.003	0.49	0.56	0.44	0.51	0.45	0.45	0.29	0.23

Table 3

Correlations of the financial instrumental variables for quarterly, annual, and monthly observations.

CBPREM is the annualized yield-to-maturity of corporate bonds rated Baa by Moody's Investor Services, minus the yield of Aaa-rated bonds. *TB1MO* is the continuously compounded rate of return of a one-month Treasury bill. *VWYLD* is the dividend yield on the CRSP value-weighted index, measured as the previous twelve months' dividend payments divided by the level of the index. *DIVDIFF* is the difference between the dividend yield of the CRSP equally-weighted index and that of the value-weighted index. *LSLOPE* is the Aaa corporate bond yield, minus the one-month Treasury bill rate. *SSLOPE* is the difference between the three-month and the one-month Treasury bill yield. *GIPX* is the continuously compounded annual growth rate of the index of industrial production. *PI* is the inverse of the index of prices of the smallest decile of common stocks on the New York Stock Exchange, relative to the average level for the preceding twelve months.

	CBPREM	*TB1MO*	*VWYLD*	*DIVDIFF*	*LSLOPE*	*GIPX*	*PI*	*SSLOPE*
	Part I: Quarterly data (1948.Q1–1987.Q3, 163 observations)							
CBPREM	1.00							
TB1MO	0.67	1.00						
VWYLD	0.18	0.02	1.00					
DIVDIFF	−0.48	−0.60	−0.04	1.00				
LSLOPE	0.51	−0.02	0.02	−0.29	1.00			
GIPX	−0.45	−0.09	−0.18	−0.15	−0.22	1.00		
PI	−0.06	0.05	0.11	0.45	−0.12	−0.14	1.00	
SSLOPE	0.44	0.37	0.05	−0.28	0.29	−0.17	0.03	1.00
	Part II: Annual data (1926–1986, 61 observations)							
TB1MO		1.00						
VWYLD		−0.23	1.00					
DIVDIFF		0.10	−0.21	1.00				
PI		−0.07	0.54	0.09			1.00	
SSLOPE		0.69	−0.15	0.09			−0.12	1.00
	Part III: Monthly data (December 1958–November 1986, 336 observations)							
CBPREM	1.00							
TB1MO	0.59	1.00						
VWYLD	0.68	0.74	1.00					
DIVDIFF	−0.42	−0.50	−0.56	1.00				
LSLOPE	0.49	−0.18	0.18	−0.25	1.00			
GIPX	−0.56	−0.20	−0.37	−0.01	−0.23	1.00		
PI	0.14	0.16	0.09	−0.10	−0.01	0.01	1.00	
SSLOPE	0.39	0.25	0.36	−0.23	0.30	−0.23	0.066	1.00

three-month Treasury bill return. (Deflated returns produce similar results.) These regressions suggest that the instruments should allow us to construct powerful tests of the Euler equation. The signs and magnitudes of the coefficients on the asset returns are consistent with earlier studies. For example, *VWYLD*(-1) enters with a positive coefficient and *TB1MO* enters with a negative coefficient in each regression. The R-squares (right-tail probability values) for the quarterly sample range from 0.12 (0.02) to 0.26 (0.00) across the portfolio returns. In annual regressions, the range is from 0.12 (0.29) to 0.48 (0.00). In monthly data, the range is from 0.06 (0.01) to 0.11 (0.00). The instruments seem to capture changing expected excess returns in both the bond and stock markets; they are less strongly related to the growth rates of the future consumption expenditures. The R-squares in regressions predicting the growth of nondurables consumption expenditures are 0.09, 0.12, and 0.02, respectively, in quarterly, annual, and monthly data. The corresponding right-tail probability values are 0.09, 0.29, and 0.66.

5. Empirical results

In tables 4–7 we present the results of estimating and testing the models using monthly, quarterly, and annual consumption expenditures and returns data. Since earlier work focuses on monthly data, we start with it. We use five assets: the three common stock portfolios from size deciles 1, 5, and 10, a long-term government bond portfolio, and a one-month Treasury bill.

In the first panel of table 4 we present results using seven instruments: a constant and three variables lagged once and twice in relation to when the five asset returns are realized. The three variables are real consumption growth over one month, the real return of a one-month Treasury bill, and the real one-month return of the small-stock portfolio. The errors u_{t+1} in the Euler equation $\mathrm{E}_t[u_{t+1}]=0$ are assumed to follow an MA(0) process in the time-separable model and an MA(1) process in the one-lag model.

In the first row of table 4 the notation $b_1 \equiv 0$ states that the nonseparability parameter is set equal to zero and we are then estimating and testing the time-separable model. The point estimate of the subjective discount rate in this case is 0.993 and the estimate of the concavity parameter A is 0.31. The right tail p-value for the goodness-of-fit test is 0.006, however, and the model is rejected. Rejecting this model is consistent with the earlier conclusions of Hansen and Singleton (1982) and others.

In the second row of table 4 the nonseparability parameter b_1 is estimated along with the parameters β and A. The model is not rejected by the goodness-of-fit test, the p-value being 0.15. The subjective discount rate is close to one and the concavity parameter is $A=2.1$. As we argue in the appendix, in the nonseparable model the concavity parameter is approximately equal to the RRA coefficient but may differ substantially from the inverse of the elasticity of consumption with respect to the interest rate. The point estimate of the nonseparability parameter b_1 is positive and significantly different from zero. Taking the result at face value, it provides evidence that durability of consumption expenditures dominates habit persistence in a one-lag model, consistent with the earlier findings of Dunn

Table 4

Test of models with habit persistence or durability of consumption expenditures using monthly returns data for May 1959–October 1986, 330 observations, and monthly consumption.

The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is consumption expenditures at date t. A is the concavity parameter, β is the rate-of-time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). In the time-separable model, b_1 is set equal to zero. Estimation is by generalized method of moments (GMM). Asymptotic standard errors are in parentheses. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the GMM criterion function. The tests use a system of five asset returns: the common stock portfolios from size deciles 1, 5, and 10 (small, medium, and large), a long-term government bond, and a one-month Treasury bill. Real returns are the nominal returns deflated by the price deflator corresponding to the measure of consumption.

Instruments[a]	Consumption[b]	β	A	b_1	χ^2	p-value
	Panel 1: Using the most recent lags of the instruments[c]					
Cons. & rets. (lags 1 and 2)	Nondurables	0.993 (0.000)	0.305 (0.073)	$\equiv 0$[e]	57.23	0.006
		1.001 (0.001)	2.112 (0.531)	0.427 (0.065)	40.25	0.150
Financial (lag 1)	Nondurables	0.999 (0.000)	−0.046 (0.186)	$\equiv 0$[e]	77.87	0.001
		0.838 (0.062)	8.437 (10.397)	−0.717 (0.194)	45.02	0.347

and Singleton (1986), Eichenbaum, Hansen, and Singleton (1988), and Eichenbaum and Hansen (1990). This result, however, is not robust.

In the lower portion of panel 1 of table 4 we replace the lagged consumption and return instruments with a constant and the lagged financial variables summarized in table 1. The point estimate of b_1 is now -0.72 and is significantly different from zero. The p-value of the goodness-of-fit test is high, 0.347. These numbers suggest that habit persistence dominates durability of consumption expenditures.[2]

These results serve two purposes. First, they demonstrate that we can replicate earlier results using our data sample. Second, they provide a warning that the estimation of the nonseparability parameter is potentially sensitive to the choice of instruments.

[2]The GMM estimate of b_1 is based on a marginal utility that is infinite for $c_t + b_1 c_{t-1} = 0$ and undefined for $c_t + b_1 c_{t-1} < 0$. Our estimates of b_1 ensure that the argument of the utility function, $c_t + b_1 c_{t-1}$, is positive at all dates. In our quarterly sample the minimum ratio (c_t/c_{t-1}) is 0.98 and in the annual and monthly samples it is 0.92 and 0.98, verifying that the argument of the utility function is positive at all dates and for all point estimates of b_1 presented in tables 4–7.

Table 4 (continued)

Instruments[a]	Consumption[b]	β	A	b_1	χ^2	p-value
	Panel 2: Using the second lagged values of the instruments[c]					
Cons. & rets. (lag 2)	Nondurables	0.999 (0.001)	1.105 (0.408)	$\equiv 0$[e]	21.26	0.266
		0.999 (0.001)	1.154 (0.565)	0.065 (0.227)	18.50	0.358
Financial (lag 2)	Nondurables	0.999 (0.000)	0.657 (0.239)	$\equiv 0$[e]	75.14	0.002
		0.837 (0.060)	12.152 (13.538)	−0.642 (0.215)	55.85	0.075
Panel 3: Using the second lagged values of the instruments and a higher-order moving-average process for the error terms[d]						
Cons. & rets. (lag 2)	Nondurables	0.999 (0.001)	1.020 (0.381)	$\equiv 0$[e]	18.32	0.435
		0.999 (0.001)	1.204 (0.585)	0.083 (0.228)	16.76	0.471
Financial (lag 2)	Nondurables	0.999 (0.001)	1.113 (0.291)	$\equiv 0$[e]	63.20	0.024
		0.999 (0.002)	1.918 (1.753)	−0.361 (0.223)	47.60	0.255

[a]The financial instruments consist of a constant and the eight variables summarized in table 1. The notation '(lags 1 and 2)' indicates that the variables are lagged one month and two months in relation to the asset returns in the Euler equations. When the financial instruments are denoted '(lag 2)', they are lagged two months. 'Cons. & rets.' denotes an instrument set composed of: a constant, the growth of the consumption measure, the real Treasury bill return, and the real return of the size portfolio from the smallest decile of firms. When denoted 'Cons. & rets. (lags 1 and 2)', each of the variables is lagged one period and two periods in relation to the asset returns in the Euler equation, and there are seven instruments. When the instruments are denoted 'Cons. & rets. (lag 2)', the consumption and returns are lagged two periods only and there are four instruments.

[b]Monthly consumption data are real per-capita consumer expenditures for nondurable goods.

[c]In panels 1 and 2, the error terms are assumed to follow an MA(0) process when the time-separable model ($b_1 \equiv 0$) is estimated and an MA(1) process when the one-lag model is estimated.

[d]In panel 3, the error terms are assumed to follow an MA(1) process when the time-separable model ($b_1 \equiv 0$) is estimated and an MA(2) process when the one-lag model is estimated.

[e]An '$\equiv 0$' indicates that the parameter is set to zero.

In panel 2 of table 4 we exclude the first lagged values of the instruments. This experiment can be motivated by concerns about measurement errors in the data, publication lag, or time aggregation. In the first subpanel, with lagged consumption and returns as the instruments, the point estimate of b_1 is 0.065 and the standard error is 0.227, providing no evidence that either habit persistence or durability plays a dominant role. Moreover, the tests do not reject the time-separable model. In the lower part of panel 2, using the second lags of the financial variables, we estimate a negative value of b_1 that

Table 5

Test of models with habit persistence or durability of consumption expenditures using quarterly returns data for 1948.Q2–1986.Q2, 153 observations.

The model assumes that a representative agent maximizes

$$E_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is real per-capita nondurables consumption expenditures at date t. The consumption expenditures are seasonally adjusted by the Commerce Department using the X-11 seasonal adjustment program. A is the concavity parameter, β is the rate of time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). Estimation is by generalized method of moments (GMM). The error terms are assumed to follow an MA(0) process when the time-separable model ($b_1 \equiv 0$) is estimated and an MA(1) process when the one-lag model is estimated. Asymptotic standard errors are in parentheses. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the GMM criterion function. Dec$N$ is the real return of common stocks from market-value decile N. The real returns are the nominal returns deflated by the nondurables price deflator.

System	β	A	b_1	χ^2	p-value
Panel 1: Using the financial instruments' most recent lagged values[a]					
Treasury bill	1.042	6.31	$\equiv 0$[c]	61.43	0.034
Government bond	(0.009)	(1.31)			
Stocks: Dec1					
Stocks: Dec5	0.883	0.78	−0.95	37.08	0.686
Stocks: Dec10	(0.049)	(1.19)	(0.05)		
Treasury bill	0.995	−0.36	$\equiv 0$[c]	38.69	0.001
Stocks: Dec10	(0.003)	(0.40)			
	0.575	4.94	−0.95	22.36	0.099
	(0.162)	(8.71)	(0.08)		

is nearly three standard errors from zero. The time-separable model is strongly rejected by the goodness-of-fit test.

In panel 3 we repeat the experiment of panel 2, except that we increase the order of the moving-average process assumed for the error terms. In the time-separable model ($b_1 \equiv 0$) we use an MA(1) assumption and in the one-lag model we use an MA(2). The results are similar to those of panel 2. When lagged consumption and returns are the instruments, the estimate of b_1 is close to zero and the time-separable model is not rejected. When the lagged financial variables are the instruments, the time-separable model is rejected. The point estimate of b_1 is negative and the time-nonseparable model is not rejected.

We conclude from table 4 that the evidence of durability in the monthly consumption expenditures relies on using the most recent lagged values of the endogenous variables as instruments. There are reasons to be suspicious of results that rely on these instruments, however. The first lagged values of consumption and returns are suspect, given the potential problems with measurement errors, publication lag, and time aggregation. In contrast,

Table 5 (continued)

System	β	A	b_1	χ^2	p-value
Panel 2: Using the financial instruments lagged two periods[a]					
Treasury bill	1.002	2.82	$\equiv 0$[c]	65.14	0.016
Government bond	(0.002)	(0.35)			
Stocks: Dec1					
Stocks: Dec5	0.998	1.81	−0.97	34.39	0.792
Stocks: Dec10	(0.002)	(0.83)	(0.01)		
Treasury bill	0.998	2.09	$\equiv 0$[c]	46.85	0.000
Stocks: Dec10	(0.002)	(0.37)			
	0.930	1.82	−0.97	13.45	0.568
	(0.082)	(1.59)	(0.01)		
Panel 3: Using lagged consumption and returns as instruments[b]					
Treasury bill	0.999	0.63	$\equiv 0$[c]	48.92	0.037
Government bond	(0.001)	(0.23)			
Stocks: Dec1					
Stocks: Dec5	0.998	0.67	−0.28	34.95	0.329
Stocks: Dec10	(0.001)	(0.29)	(0.10)		
Treasury bill	0.999	0.71	$\equiv 0$[c]	29.49	0.003
Stocks: Dec10	(0.001)	(0.28)			
	0.999	1.70	−0.21	19.98	0.046
	(0.005)	(1.08)	(0.15)		

[a]The financial instrumental variables are a constant and the eight variables summarized in table 1.

[b]The instruments are a constant, the growth of the consumption measure, the real Treasury bill return, and the real return of the size portfolio from the smallest decile of firms; each variable is lagged one period and two periods in relation to the asset returns in the Euler equation.

[c]An '$\equiv 0$' indicates that the parameter is set to zero.

whether we use the first or the second lags of the financial variables, we estimate negative b_1 coefficients, which suggests that habit persistence dominates durability. Of course, given the problems with monthly consumption data, evidence based on the monthly data is not conclusive.

In tables 5 and 6 we present results using quarterly and annual consumption expenditures and returns. We use as instruments either the financial variables lagged by one period, the financial variables lagged by two periods, or the consumption and return variables lagged by one and two periods. Table 5 presents quarterly data, assuming that decisions are made quarterly. In table 6 we assume an annual decision interval and use annual data.

In table 5, the results use two systems of asset returns. The first is similar to the monthly five-asset system and consists of the three stock portfolios (deciles 1, 5, and 10), a portfolio of long-term government bonds, and a strategy of rolling over one-month Treasury bills. In this system the model is

Table 6

Test of models with habit persistence or durability of consumption expenditures using annual returns for 1932–1984, 53 observations, and annual consumption.

The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is real per-capita nondurables consumption expenditures at date t. A is the concavity parameter, β is the rate-of-time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). Estimation is by generalized method of moments (GMM). The error terms are assumed to follow an MA(0) process when the time-separable model ($b_1 \equiv 0$) is estimated and an MA(1) process when the one-lag model is estimated. Standard errors are in parentheses. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the GMM criterion function. Dec$N$ is the real return of common stocks from market-value decile N. The real returns are the nominal returns deflated by the nondurables price deflator.

System	β	A	b_1	χ^2	p-value
Panel 1: Using the financial instruments' most recent lagged values[a]					
Treasury bill	1.015	0.85	$\equiv 0$[c]	35.76	0.149
Government bond	(0.004)	(0.35)			
Stocks: Dec1					
Stocks: Dec5	1.167	0.03	−0.85	21.01	0.786
Stocks: Dec10	(0.021)	(0.52)	(0.16)		
Treasury bill	0.966	−2.24	$\equiv 0$[c]	19.67	0.033
Stocks: Dec10	(0.008)	(0.34)			
	0.995	0.05	−0.79	16.62	0.055
	(0.014)	(0.55)	(1.16)		

challenged to explain the differences in the returns of common stocks grouped by firm size, which have served as an acid test of the capital asset pricing model. The second system, consisting of just two assets, the Treasury bills and the large stocks, directs attention to the differences in the return of Treasury bills and stocks.

In panel 1 of table 5 the instruments are a constant and the eight financial variables lagged once. In panel 2 the financial instruments at lag two are used. The point estimates of the parameter b_1 all lie between -0.95 and -0.97 and are many standard errors away from zero. The goodness-of-fit tests show a markedly improved fit when the nonseparability parameter is included. The results suggest that the effects of habit persistence dominate the effects of durability in the quarterly data.

We replicated panel 2 of table 5, except that we use an MA(1) assumption for the error terms in the time-separable model and an MA(2) assumption in the one-lag, nonseparable model. The point estimates of the parameters are similar but the standard errors are, in most of the cases, larger. The goodness-of-fit statistics are typically smaller than in panel 2 of table 5. For

Table 6 (continued)

System	β	A	b_1	χ^2	p-value
Panel 2: Using the financial instruments lagged two periods[a]					
Treasury bill Government bond Stocks: Dec1	0.966 (0.007)	−3.09 (0.59)	$\equiv 0$[c]	34.17	0.195
Stocks: Dec5 Stocks: Dec10	0.978 (0.001)	0.45 (0.02)	−0.92 (0.01)	20.50	0.809
Treasury bill Stocks: Dec10	0.947 (0.017)	−3.70 (0.97)	$\equiv 0$	19.29	0.037
	0.994 (0.099)	−3.30 (4.19)	−0.62 (0.18)	8.47	0.487
Panel 3: Using lagged consumption and returns as instruments[b]					
Treasury bill Government bond Stocks: Dec1	0.999 (0.005)	0.54 (0.32)	$\equiv 0$[c]	35.06	0.371
Stocks: Dec5 Stocks: Dec10	1.025 (0.004)	2.58 (0.37)	−0.12 (0.06)	23.55	0.860
Treasury bill Stocks: Dec10	0.997 (0.010)	−1.09 (0.44)	$\equiv 0$[c]	26.48	0.009
	0.988 (0.007)	0.07 (0.40)	−0.74 (0.64)	14.05	0.230

[a]The financial instrumental variables are a constant and the eight variables summarized in table 1.
[b]The instruments are a constant, the growth of the consumption measure, the real Treasury bill return, and the real return of the size portfolio from the smallest decile of firms; each variable is lagged one period and two periods in relation to the asset returns in the Euler equation.
[c]An '$\equiv 0$' indicates that the parameter is set to zero.

example, the time-separable model is no longer rejected in the five-asset system (the p-value is 0.44) but is rejected in the two-asset system (the p-value is 0.006). The nonseparable model is not rejected. Including additional moving-average terms when the autocorrelations are not significant implies that the estimate of the covariance matrix of the orthogonality conditions will be noisier. We would expect this to reduce the efficiency of the estimates and the power of the goodness-of-fit tests.

In the monthly data, evidence of durability in consumption expenditures emerged only when the first lagged values of the consumption and returns were used as instruments. In panel 3 of table 5 we replicate this experiment with quarterly data. The results show that, merely by switching from monthly to quarterly data, we lose the evidence in favor of durability. For the five-asset system the coefficient b_1 is negative and statistically different from zero and for the two-asset system the point estimate of b_1 is negative but insignificantly different from zero. We repeated the tests in table 5, combining the instruments from panels 2 and 3. The results are similar to those reported in panel 2.

In table 6 we repeat the tests reported in table 5 with annual rather than quarterly data. These results provide further evidence that habit persistence dominates durability. All of the point estimates of b_1 are negative and introducing nonseparability improves the goodness-of-fit. The annual data, of course cover a longer sample period and may be less influenced by measurement errors. Furthermore, results using the annual data should not be affected by problems with the seasonality of consumption expenditures and returns within the year.

We replicated the tests in panel 2 of table 6 using the financial variables at lag two as instruments and using the alternative moving-average assumptions for the error terms [MA(1) for the time-separable and MA(2) for the one-lag model]. These experiments show that the results are robust. In the five-asset system the point estimate of b_1 is -0.87 (standard error 0.03). None of the models can be rejected by the goodness-of-fit statistics, however.

Our empirical results in tables 4–6 may be understood in terms of the time-series properties of consumption and the predictive ability of the different instruments for asset returns and consumption. In monthly data, consumption growth rates are negatively autocorrelated. In quarterly data the autocorrelation is closer to zero, and in annual data the autocorrelation is positive (see table 2). Recall that durability of consumption expenditures induces negative autocorrelation. Habit persistence induces positive autocorrelation in consumption growth, since the consumer optimally smoothes consumption more than with time-separable preferences. When we use lagged consumption and returns as the instruments, the autocorrelation properties of consumption are given more weight in the Euler restrictions. Therefore, when we use monthly consumption data and lagged consumption instruments, we find evidence for durability in monthly data.

When the lagged financial instruments are used in the Euler equation, the predictive power is relatively higher for the asset returns, and lower for future consumption, than when lagged consumption growth rates and returns are used. Habit persistence tends to reduce the elasticity of consumption with respect to investment returns (see the appendix), which suggests relatively higher volatility in expected returns in comparison with consumption changes. These effects receive more weight in the Euler restrictions when the lagged financial variables are the instruments. We therefore find evidence for habit persistence, even in the monthly data, when the lagged financial variables are used.

Fig. 1 illustrates the sensitivity of the one-lag model to the value of the nonseparability parameter b_1. The values of the objective function are plotted, minimized over the choice of A and β, for given values of b_1. An MA(1) weighting matrix is used. Results for annual data and five assets are plotted. (Similar patterns are observed in the other cases.) The objective

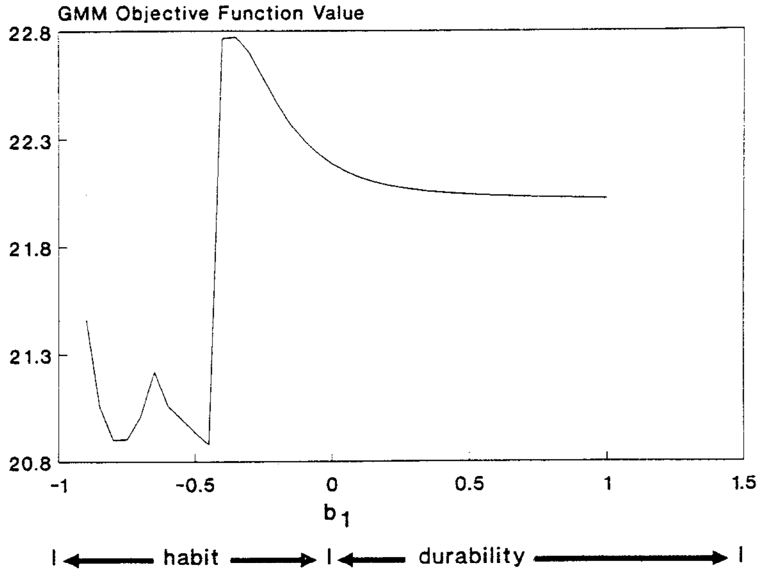

Fig. 1. Sensitivity of the GMM objective function to the nonseparability parameter.

The y-axis is the value of the generalized-method-of-moments objective function, minimized over the choice of the parameters A and β, for given values of b_1, the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). The b_1 values are shown on the x-axis. The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is consumption expenditures at date t.

The consumption data are annual expenditures for consumer nondurable goods. The real asset returns are common stocks from size deciles 1, 5, and 10, a long-term government bond, and the return to rolling over one-month Treasury bills for 1932–1984, 53 observations. The instruments are a constant and the five financial variables described in the text.

function is highly nonlinear in the parameter b_1. Typically, we find that there is a local minimum in the region of durability ($b_1 > 0$), and a 'hill' in the objective function over which the algorithm must climb to attain the global minimum in the habit-persistence region ($b_1 < 0$).

Table 7 reports tests of the hypothesis H_0', that the parameter b_1 is zero, allowing the error term to be MA(1), against the alternative that b_1 is not zero. Two test statistics are reported. The two statistics have the same asymptotic distribution under the null hypothesis. One is the square of the t-ratio for the b_1 parameter. The second test uses the approach described in Eichenbaum, Hansen, and Singleton (1988). The restricted system (imposing $b_1 = 0$) is estimated using the GMM weighting matrix from the unrestricted system, allowing the MA(1) error structure. The difference between quadratic

Table 7

Tests for the importance of a nonzero nonseparability parameter with a moving-average error structure. The quarterly returns data are for 1948.Q2–1986.Q2, 153 observations, and the annual returns data are for 1932–1984, 53 observations.

The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is nondurables consumption expenditures at date t. A is the concavity parameter, β is the rate of time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). Estimation is by generalized method of moments (GMM), using a weighting matrix that assumes the error terms are autocorrelated at lag 1 and uncorrelated at longer lags. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the test statistic. The $\chi^2(1)$ statistic is the difference between the GMM criterion functions, imposing the hypothesis that b_1 is zero and leaving b_1 unrestricted. The $t(b_1)^2$ statistic is the square of the t-statistic of b_1 from tables 5 and 6.

System	$\chi^2(1)$	p-value	$t(b_1)^2$	p-value
		Quarterly data		
Five assets	23.86	0.00	329.9	0.00
Two assets	4.06	0.04	126.1	0.00
		Annual data		
Five assets	9.75	0.00	28.22	0.00
Two assets	1.90	0.17	0.46	0.49

forms using the restricted and the unrestricted objective function is asymptotically distributed as a chi-square variate with one degree of freedom.

Table 7 shows that the two statistics imply the same inferences, so there is no signal of departures from the large-sample properties of the statistics. The tests reject the null hypothesis H_0' against H_a in three of four cases. In the fourth case, the two-asset system with annual data, neither test can reject the null. Overall, the evidence supports a negative value of the nonseparability parameter.

We find that direct estimation of the two-lag nonseparable model is problematic. It is difficult to estimate both b_1 and b_2 with precision. Fig. 2 illustrates the problem. We establish a grid of values for b_1 and b_2 in increments of 0.1. At each point in the grid we condition on these values and search for A and β to minimize the objective function. Fig. 2 shows the right-tail probability value associated with the minimized value of the objective function on the vertical axis, plotted against the specified values of $(-b_1)$ on the x-axis and $(-b_2)$ on the relief axis. The figure therefore represents the model's goodness-of-fit for particular values of the nonseparability parameters. An MA(2) weighting matrix is used for every point on the grid.

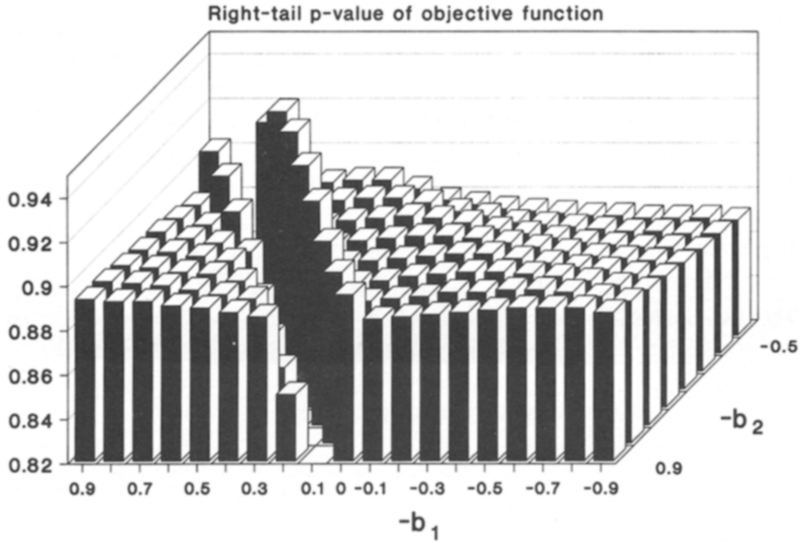

Fig. 2. Sensitivity of the GMM objective function to the nonseparability parameters in a two-lag model.

The model assumes that a representative agent maximizes

$$E_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1} + b_2 c_{t-2}$ and c_t is consumption expenditures at date t. A is the concavity parameter, β is the rate of time discount, and b_1 and b_2 are the parameters representing habit persistence or durability. The vertical axis is the right-tail probability value for the generalized-method-of-moments objective function, minimized over the choice of the parameters A and β, for fixed values of b_1 and b_2. The negatives of the b-values are shown on the other two axes. Negative values of the b's (positive numbers on the axes) indicate habit persistence, and positive values of the b's indicate durability of goods.

The quarterly real asset returns are for common stocks from size deciles 1, 5, and 10, a long-term government bond, and rolling over one-month Treasury bills for 1948.Q2–1986.Q2, 153 observations. The real returns are the nominal returns deflated by the nondurables price deflator. The consumption data are real per-capita expenditures for consumer nondurable goods, seasonally adjusted by the Commerce Department using the X-11 program. The instruments are a constant and the eight financial variables summarized in table 1.

Results for the five-asset system and quarterly data are shown. A point in the center of the x-relief plane corresponds to the time-separable model ($b_1 = 0$, $b_2 = 0$). Moving toward the right rear corner implies increasing durability and moving toward the left front corner implies habit persistence. There is a ridge, along which b_1 and b_2 add up to about -0.90, in the habit-persistence region of the figure. The value of the objective function is insensitive to the individual coefficients b_1 and b_2. Therefore, the individual coefficients cannot be estimated reliably. We find a similar result when we estimate the two-lag model with annual data or with quarterly durable goods expenditures, or when we use a constant as the only instrumental variable.

Although our results suggest that habit persistence is empirically relevant, the two separate coefficients cannot be estimated reliably in the two-lag model. Therefore, we cannot infer the half-life of habit persistence. Even in the absence of durability, two parameters (h and a) determine the half-life of habit persistence. Three parameters would be needed to isolate the separate effects of durability and habit persistence [see eq. (5)], so we are unable to isolate these effects.

6. The robustness of the empirical results

6.1. Seasonal adjustment of consumption

The estimates and test results reported in tables 4–7 are based on seasonally adjusted data, as explained in section 4. The nonseparable models imply that the flow of services and the subsistence level are averages of lagged consumption expenditures. But seasonally adjusted consumption data have already been smoothed with the X-11 seasonal adjustment program. The smoothing could bias the tests and the parameter estimates.

The similarity of our findings using quarterly and annual data suggests that the results are not driven by seasonality or by the particular seasonal adjustment used within the year. An appealing way to verify further that the results are not sensitive to seasonal adjustment is to use unadjusted quarterly expenditures data and to allow for seasonal variation in the utility function as in Miron (1986), Ferson and Harvey (1991), and English, Miron, and Wilcox (1989). These studies introduce taste-shift parameters that allow the utility for a given level of consumption to vary with the season. We attempted in some experiments to incorporate seasonal taste-shift parameters together with the nonseparability parameter b_1 in the model, but were unable to estimate a model reliably with the additional parameters. We therefore report the results of a less ambitious experiment.

We adopt a two-step approach. In the first step, we regress the logarithm of unadjusted quarterly real per-capita consumption expenditures on a time trend and dummy variables indicating the quarter. We take the residuals from these regressions, add back the sample means, and exponentiate. The resulting consumption series is seasonally adjusted, but avoids the moving averages used by X-11. The procedure is similar to assuming multiplicative seasonal taste-shift parameters, as in Miron (1986). In the second step, we use the dummy adjusted data to estimate and test the Euler equations (8) and (9). In principle, one would like to combine the two steps, but the

Table 8

Tests of models with habit persistence or durability of consumption expenditures using quarterly data for 1948.Q2–1986.Q2, 153 observations, and seasonal adjustment via dummy variables.

The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is consumption expenditures for nondurable goods at date t. A is the concavity parameter, β is the rate-of-time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). Estimation is by generalized method of moments (GMM). Asymptotic standard errors are in parentheses. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the GMM criterion function. The instrumental variables are a constant and the eight variables summarized in table 1. Dec$N$ is the real return of common stocks from market-value decile N. The real returns are the nominal returns deflated by the nondurables price deflator.

System	β	A	b_1	χ^2	p-value
Treasury bill	0.997	0.441	$\equiv 0$[a]	57.97	0.063
Government bond	(0.001)	(0.235)			
Stocks: Dec1					
Stocks: Dec5	0.926	1.724	−0.791	39.54	0.579
Stocks: Dec10	(0.061)	(1.833)	(0.113)		
Treasury bill	0.998	0.213	$\equiv 0$[a]	37.89	0.001
Stocks: Dec10	(0.001)	(0.238)			
	0.946	3.104	−0.621	23.31	0.078
	(0.044)	(4.496)	(0.261)		

[a]An '$\equiv 0$' indicates that the parameter is set to zero.

two-step approach does indicate the potential sensitivity of our findings to the method of seasonal adjustment.[3]

The results are summarized in table 8 and are similar to those in table 5. The goodness-of-fit statistics imply that habit persistence improves the fit of the model. The point estimates of the nonseparability parameter b_1 are negative and several standard errors from zero.

6.2. Expenditures on durable goods

The Commerce Department provides data on three groups of consumer expenditures, which are labeled durables, nondurables, and services. Our

[3]The seasonally adjusted consumption levels are given by $c_t \exp(\gamma'[\bar{D} - D_t])$, where c_t is the unadjusted consumption, γ is a vector of three shift parameters relative to the first quarter, D_t is a vector of dummy variables for quarters 2–4, and $\bar{D}$ is the vector of sample means of the dummy variables. Miron (1986) includes both time and a time-squared term in his regressions for the log of consumption. In a previous version of this paper we followed Miron by including the squared term. Our results using this alternative series were very similar to those reported in tables 8, 9, and 10.

estimates and tests reported in sections 5 and 6.1 used the nondurables series only. We recognized that the Commerce Department's nondurables series may in fact be durable, and we accounted for the durability in the derivation of the nonseparable models. Our approach is formally justified under the assumption that the consumer's preferences are separable over the flow of services from the Commerce Department's durables, nondurables, and services.

Dunn and Singleton (1986), Eichenbaum, Hansen, and Singleton (1988), and Eichenbaum and Hansen (1990) derive, estimate, and test Euler equations for models with nonseparable preferences across goods. Parameters are set for the degree of substitution across the goods. If we allow for such preferences, which are both time-nonseparable and nonseparable across goods, the number of parameters in the model increases.

In this section we adopt the hypothesis that the preferences are separable across goods and focus on the Euler equation implied by time-nonseparable preferences, defined over the flow of services from the Commerce Department's durable goods expenditure series. Results for annual and quarterly data (both X-11 and dummy seasonally adjusted versions) are reported in table 9. One might expect b_1 to be positive for consumer durables, but the point estimates of b_1 – although greater than the estimates for nondurables – are less than zero in all six cases. The goodness-of-fit tests indicate an improved fit with the negative b_1 coefficients. We conclude that our earlier finding that habit persistence dominates durability is not due to our exclusion of the consumption flows from the durable goods expenditure series.

Fig. 3 illustrates the sensitivity of the objective function to the value of the nonseparability parameter b_1, using durable-goods expenditures. The values of the objective function, minimized over the choice of A and β, are plotted for fixed values of b_1. Results for quarterly (X-11 seasonally adjusted) data and all five assets are plotted. Similar patterns are observed in the other cases, and the conclusion is similar to that obtained from fig. 1. Typically, however, a local minimum in the region of durability is closer to the global minimum in the region of habit persistence when durable-goods expenditures are used than when nondurables are used.

Since the results for nondurable and durable goods are similar, the sum of the two is unlikely to lead to different conclusions. There are potential difficulties, however, in the interpretation of the experiments in table 9 and fig. 3. If preferences are not separable across durable and nondurable goods, then applying the model to data on durable goods creates a missing-variables problem. Because nondurable- and durable-goods expenditures are correlated, the durable expenditures may proxy for the missing nondurables. The estimates of the nonseparability parameter may thus be biased toward habit persistence. A multi-good model could potentially control for this effect,

Table 9

Tests of models with habit persistence or durability of consumption expenditures using consumer durable goods expenditures.

The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is consumption expenditures at date t. A is the concavity parameter, β is the rate-of-time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). Estimation is by generalized method of moments (GMM). Asymptotic standard errors are in parentheses. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the GMM criterion function. The instrumental variables are a constant and the eight variables summarized in table 1. Dec$N$ is the real return of common stocks from market-value decile N.

System	β	A	b_1	χ^2	p-value
	X-11 adjusted quarterly data (1948.Q2–1986.Q2, 153 observations)				
Treasury bill	0.995	0.042	$\equiv 0$[a]	63.18	0.024
Government bond	(0.006)	(0.037)			
Stocks: Dec1					
Stocks: Dec5	0.996	0.023	−0.772	42.55	0.447
Stocks: Dec10	(0.002)	(0.084)	(0.338)		
Treasury bill	0.994	−0.042	$\equiv 0$[a]	42.61	0.003
Stocks: Dec10	(0.009)	(0.058)			
	1.274	1.011	−0.652	25.507	0.044
	(5.906)	(0.088)	(0.639)		
	Dummy adjusted quarterly data (1948.Q2–1986.Q2, 153 observations)				
Treasury bill	0.995	0.079	$\equiv 0$[a]	62.85	0.026
Government bond	(0.001)	(0.035)			
Stocks: Dec1					
Stocks: Dec5	0.953	1.307	−0.700	41.93	0.474
Stocks: Dec10	(0.059)	(1.096)	(0.102)		
Treasury bill	0.995	−0.025	$\equiv 0$[a]	42.87	0.0003
Stocks: Dec10	(0.001)	(0.052)			
	0.845	2.439	−0.620	22.52	0.095
	(0.085)	(3.130)	(0.240)		
	Annual data (1932–1984, 53 observations)				
Treasury bill	1.085	1.211	$\equiv 0$[a]	34.99	0.170
Government bond	(0.011)	(0.284)			
Stocks: Dec1					
Stocks: Dec5	1.011	0.333	−0.442	21.96	0.739
Stocks: Dec10	(0.018)	(0.589)	(0.366)		
Treasury bill	0.964	−0.484	$\equiv 0$[a]	26.16	0.004
Stocks: Dec10	(0.007)	(0.134)			
	1.016	1.075	−0.225	15.56	0.077
	(0.077)	(2.467)	(0.703)		

[a]An '$\equiv 0$' indicates that the parameter is set to zero.

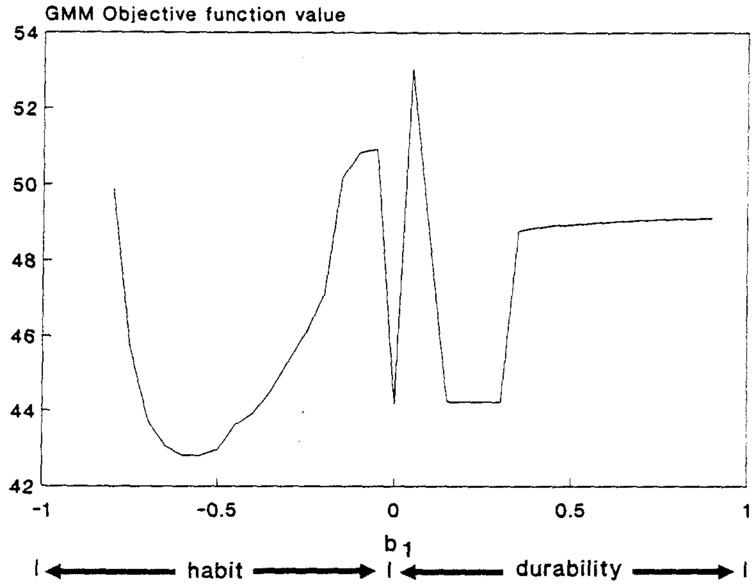

Fig. 3. Sensitivity of the GMM objective function to the nonseparability parameter in a one-lag model with durable-goods expenditures.

The y-axis is the value of the generalized-method-of-moments objective function, minimized over the choice of the parameters A and β, for given values of b_1, the parameter representing habit persistence or durability. The b_1 values are shown on the x-axis. Negative values indicate habit persistence, and positive values indicate durability of goods. The model assumes that a representative agent maximizes

$$E_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is consumption expenditures at date t.

The consumption data are quarterly expenditures for consumer durable goods, seasonally adjusted by the Commerce Department using the X-11 program. The real asset returns are for common stocks from size deciles 1, 5, and 10, a long-term government bond, and rolling over one-month Treasury bills for 1948.Q2–1986.Q2, 153 observations. The real returns are the nominal returns deflated by the consumer durables price deflator. The instruments are a constant and the eight financial variables summarized in table 1.

although the larger number of parameters will create econometric difficulties. We leave for future research a complete study of the interaction between habit persistence and the complementarity or substitutability of different consumption goods.

7. The equity premium revisited

Mehra and Prescott (1985) introduce the equity premium puzzle. They consider a pure exchange economy in which the representative consumer has time- and state-separable preferences with a constant RRA coefficient. They assume that consumption growth is a two-state Markov process and calibrate this process to match the sample mean, variance, and first-order autocorrelation of the annual growth rate of per-capita consumption in the years 1889–1978. They are unable to find a plausible pair of the subjective discount rate and the RRA coefficient to match the sample mean of the annual real rate of interest and of the equity return over the same 90-year period.

Mehra and Prescott's calibration exercise is not directly comparable to the estimation and testing of Euler equations that we report in the earlier sections. Our starting point was a system of Euler equations:

$$\mathrm{E}_t\left[R_{i(t+1)} MRS_{t+1}\right] = 1, \tag{10}$$

where $R_{i(t+1)}$ is the one-plus-return on asset i over $[t, t+1]$ and MRS_{t+1} is the marginal rate of substitution between dates and t and $t+1$. The Euler restrictions we tested included in particular the unconditional Euler equation:

$$\mathrm{E}\left[R_{i(t+1)} MRS_{t+1}\right] = 1. \tag{11}$$

Using predetermined variables z_{jt} as instruments, we included additional implications of the Euler equation, specifically:

$$\mathrm{E}\left[R_{i(t+1)} MRS_{t+1} \mid z_{jt}\right] = 1. \tag{12}$$

Including the predetermined instruments, our estimates of the *RRA* coefficient were reasonable in magnitude, for both the time-separable and the nonseparable model. In contrast, Mehra and Prescott find that a large *RRA* coefficient is implied by the mean equity premium.

Hansen and Singleton (1983) and Ferson and Harvey (1991) observe that, when the Euler equation with $b_1 = 0$ is estimated using only unconditional moments, large but imprecise estimates of the risk-aversion coefficient are found. In table 10 we repeat the tests of tables 5, 6, 8, and 9 using a constant as the only instrument. We examine the five-asset system, because the two-asset systems are not overidentified. The estimates of the concavity parameter A are typically larger in absolute magnitude than in the previous tables. When we use nondurable goods, the estimates of b_1 are negative in two of the three experiments, which suggests habit persistence. When durable-goods expenditures are used, all three point estimates are positive, which suggests durability in expenditures. None of the estimates in table 10 are precise, however, and the goodness-of-fit statistics do not indicate an improved fit when we allow for nonseparable preferences. We also replicate

Table 10

Test of models with habit persistence or durability of consumption expenditures using unconditional moment restrictions.

The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is consumption expenditures at date t. A is the concavity parameter, β is the rate-of-time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). Estimation is by generalized method of moments (GMM). Asymptotic standard errors are in parentheses. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the GMM criterion function. The instrumental variable is a constant vector of ones. Dec$N$ is the real return of common stocks from market-value decile N. The real returns are the nominal returns deflated by the price deflator for the measure of consumption.

System	β	A	b_1	χ^2	p-value
		Panel 1: Consumer nondurable goods			
		X-11 adjusted quarterly data (1948.Q2–1986.Q2, 153 observations)[a]			
Treasury bill	1.018	79.40	$\equiv 0$[c]	3.95	0.267
Government bond	(0.138)	(43.72)			
Stocks: Dec1					
Stocks: Dec5	0.461	7.44	−0.903	3.50	0.174
Stocks: Dec10	(0.762)	(87.09)	(0.906)		
		Dummy adjusted quarterly data (1948.Q2–1986.Q2, 153 observations)[b]			
Treasury bill	0.951	13.865	$\equiv 0$[c]	12.53	0.006
Government bond	(0.131)	(20.590)			
Stocks: Dec1					
Stocks: Dec5	0.908	27.321	0.360	9.58	0.008
Stocks: Dec10	(0.194)	(82.393)	(4.066)		
		Annual data (1932–1984, 53 observations)			
Treasury bill	1.119	12.20	$\equiv 0$[c]	7.62	0.054
Government bond	(0.089)	(14.10)			
Stocks: Dec1					
Stocks: Dec5	0.959	18.67	−0.293	6.26	0.044
Stocks: Dec10	(0.391)	(25.61)	(0.819)		

the tests in table 7, using only a constant as the instrument, and are unable to reject the hypothesis that b_1 is zero.

In what sense then does habit persistence explain the equity premium puzzle, as claimed in Constantinides (1990)? To answer this question, consider adding a mean-preserving spread ρ_{t+1} to the asset return $R_{i(t+1)}$ as $R_{i(t+1)} + \rho_{t+1}$, where $\mathrm{E}_t[\rho_{t+1}] = 0$ and $\mathrm{E}_t[\rho_{t+1} MRS_{t+1}] = 0$. If eq. (10) holds, so does the equation

$$\mathrm{E}_t\left[\left(R_{i(t+1)} + \rho_{t+1}\right) MRS_{t+1}\right] = 1. \tag{13}$$

This implies that conditional and unconditional tests of the Euler equation

Table 10 (continued)

System	β	A	b_1	χ^2	p-value
		Panel 2: *Consumer durable goods*			
		X-11 adjusted quarterly data (1948.Q2–1986.Q2, 153 observations)[a]			
Treasury bill	1.014	7.055	$\equiv 0$[c]	11.46	0.009
Government bond	(0.038)	(8.347)			
Stocks: Dec1					
Stocks: Dec5	0.980	15.399	0.375	6.97	0.031
Stocks: Dec10	(0.085)	(8.001)	(0.784)		
		Dummy adjusted quarterly data (1948.Q2–1986.Q2, 153 observations)[b]			
Treasury bill	0.970	4.633	$\equiv 0$[c]	11.38	0.010
Government bond	(0.093)	(7.644)			
Stocks: Dec1					
Stocks: Dec5	0.895	11.797	0.352	7.36	0.025
Stocks: Dec10	(0.228)	(15.395)	(0.803)		
		Annual data (1932–1984, 53 observations)			
Treasury bill	0.531	−6.206	$\equiv 0$[c]	10.79	0.013
Government bond	(0.131)	(1.945)			
Stocks: Dec1					
Stocks: Dec5	1.039	1.461	0.476	9.73	0.008
Stocks: Dec10	(0.057)	(11.06)	(19.42)		

[a]The X-11 adjusted data are seasonally adjusted by the Commerce Department using the X-11 program.
[b]The dummy adjusted data are seasonally adjusted by the authors using dummy variables.
[c]An '$\equiv 0$' indicates that the parameter is set to zero.

do not focus on the variance of asset returns. Likewise, matching the variance of the equity premium is not one of the goals addressed by Mehra and Prescott (1985).

By contrast, Constantinides' (1990) goal is to match both the mean and the variance of the equity premium. He demonstrates that habit persistence improves on time-separable preferences. The mean equity premium is driven by the RRA coefficient rather than the elasticity of consumption. As we demonstrate in the appendix, habit persistence has a second-order effect on the RRA coefficient and, predictably, does not improve significantly on the mean equity premium. The variance of the equity premium is driven by the elasticity of consumption. As we also demonstrate in the appendix, habit persistence decreases substantially the elasticity of consumption. This implies that a small variance in consumption growth is associated with a large variance in the equity premium.

In table 11 we provide additional evidence on this issue. We repeat the tests of table 5, but replace the real return of the Treasury bill in the Euler equation with the real return plus a parameter L. This parameter allows the unconditional mean return of the bill, and therefore the mean excess return or premium of an asset relative to the bill, to be unrestricted. The L

Table 11

Test of models with habit persistence or durability of consumption expenditures using an unrestricted mean Treasury bill return, quarterly data, 1948.Q2–1986.Q2, 153 observations.

The model assumes that a representative agent maximizes

$$\mathrm{E}_0\left[(1-A)^{-1}\sum_{t=0}^{\infty}\beta^t C_t^{1-A}\right]$$

where $C_t = c_t + b_1 c_{t-1}$ and c_t is consumption expenditures at date t. A is the concavity parameter, β is the rate of time discount, and b_1 is the parameter representing habit persistence ($b_1 < 0$) or durability ($b_1 > 0$). The real Treasury bill return in the model is replaced by the real return plus the parameter, L, to leave the mean real return of the bill unrestricted. Estimation is by generalized method of moments (GMM). Asymptotic standard errors are in parentheses. P-value is the probability that a χ^2 variate exceeds the minimized sample value of the GMM criterion function. The instrumental variables are a constant and the eight variables summarized in table 1. Dec$N$ is the real return of common stocks from market-value decile N. The real returns are the nominal returns deflated by the nondurables price deflator.

System	β	A	b_1	$L(\times 100)$	χ^2	p-value
			Consumer nondurable goods (seasonally adjusted by the X-11 program)			
Treasury bill	0.997	0.13	$\equiv 0$[a]	0.07	61.23	0.028
Government bond	(0.002)	(0.22)		(0.19)		
Stocks: Dec1						
Stocks: Dec5	0.862	0.85	−0.95	0.47	36.83	0.656
Stocks: Dec10	(0.049)	(1.21)	(0.05)	(0.20)		
Treasury bill	0.982	0.04	$\equiv 0$[a]	1.61	37.91	0.001
Stocks: Dec10	(0.005)	(0.29)		(0.51)		
	0.976	0.87	−0.86	1.52	21.08	0.100
	(0.021)	(3.55)	(0.31)	(0.55)		

[a]An '$\equiv 0$' indicates that the parameter is set to zero.

parameter can be interpreted as a pricing error, similar to a 'Jensen's alpha' in the capital asset pricing model, or as a liquidity premium on the bill return that isn't captured by the model.

Table 11 shows that the point estimates of L are positive and, in the two-asset system, significantly different from zero. This implies that the average bill return that best fits the model is higher than the historical average return, consistent with the pattern in Mehra and Prescott (1985). But the goodness-of-fit tests and the estimates of b_1 do not change much from the results in table 5, so the tests are not highly sensitive to the average interest rate. This supports the interpretation that habit persistence improves the fit of the model largely through its effect on moments other than the mean equity premium. We leave for future research a more exhaustive analysis of the ability of the models to fit specific moment conditions. [See Allen (1990) for some preliminary work along these lines.]

8. Suggestions for future research

Our study suggests several avenues for future research. One extension is a multiple-good model, which may provide further evidence on substitution across consumption goods in the presence of habit persistence and durability. Such a model could potentially identify the separate effects, but our results suggest that more complex specifications will be econometrically challenging. Further work is required to understand better the relation of seasonality and habit persistence. Another challenge is to model several frequencies of consumption and asset-holding-period returns simultaneously. Such an approach is potentially promising, given the different time-series properties of asset returns and consumption over different holding periods and our evidence suggesting nonseparabilities that operate at different frequencies.

Appendix

With habit persistence and/or durability, the RRA coefficient and the elasticity of consumption with the interest rate depend on the parameters of the probability distribution of the asset returns. With plausible assumptions about the parameters of the probability distribution, Constantinides (1990) proves that the RRA coefficient approximately equals the parameter A and that the elasticity of consumption can be substantially lower than the inverse of the relative-risk-aversion coefficient. These insights are important in the interpretation of our empirical findings and are illustrated here in the context of a simple deterministic economy.

We assume that the investment opportunity set consists of just one asset, which is riskless and has (one plus) rate of return R each period. The representative consumer's preferences are the special case of the preferences represented by eqs. (3)–(4) with $\delta_0 = 1$, $b_1 = -h$, and $b_\tau = 0$ for $\tau \geq 2$. The preferences then become

$$(1-A)^{-1} \sum_{t=0}^{\infty} \beta^t (c_t - hc_{t-1})^{1-A}. \tag{A.1}$$

The consumer's initial endowment is W_0. The consumer receives zero future endowment or labor income. Therefore the equation of motion of wealth, in units of the consumption good, is

$$W_t = (W_{t-1} - c_{t-1})R. \tag{A.2}$$

An admissible consumption plan is defined by the properties $0 \leq c_t < \infty$,

$0 \leq c_t - hc_{t-1}$, and $0 \leq W_t$ for all t. The conditions

$$W_0 - [hRc_{-1}/(R-h)] \geq 0 \tag{A.3}$$

and

$$R > h \tag{A.4}$$

guarantee that the set of admissible plans is not empty. For example, the consumption plan $c_t = h^{t+1}c_{-1}$ is feasible because it implies $c_t > 0$, $c_t - hc_{t-1} = 0$, and has a discounted present value $\Sigma_{t=0}^{\infty} R^{-t} h^{t+1} c_{-1} = hRc_{-1}/(R-h)$, which is less than or equal to W_0, by the condition (A.3). Finally, the right-hand part of the inequality

$$1 < (\beta R)^{1/A} < R \tag{A.5}$$

guarantees that the utility of consumption over the infinite horizon is finite under all feasible consumption plans.

Define the derived utility of wealth at time t by the recursive equation

$$V(W_t, c_{t-1}) = \max_{c_t} \left[(1-A)^{-1}(c_t - hc_{t-1})^{1-A} + \beta V((W_t - c_t)R, c_t) \right]. \tag{A.6}$$

The first-order condition with respect to c_t is

$$(c_t - hc_{t-1})^{-A} - \beta R V_1((W_t - c_t)R, c_t) + \beta V_2((W_t - c_t)R, c_t) = 0, \tag{A.7}$$

where V_1 and V_2 are the derivatives of V with respect to its first and second arguments, respectively. The solution of (A.6) and (A.7) gives the optimal consumption plan as

$$c_t = \{R - (\beta R)^{1/A}\}(R-h)R^{-2}W_t + (\beta R)^{1/A} hR^{-1} c_{t-1} \tag{A.8}$$

and the derived utility of wealth as

$$V(W_t, c_{t-1}) = \frac{K}{1-A}\left(W_t - \frac{hRc_{t-1}}{R-h}\right)^{1-A}, \tag{A.9}$$

where

$$K \equiv \left\{R - (\beta R)^{1/A}\right\}^{-A} (R-h)^{1-A} R^{2A-1}. \tag{A.10}$$

We proceed to derive the time series of consumption. Eq. (A.8) gives W_t in terms of c_t and c_{t-1}, and the one-lag version gives W_{t-1} in terms of c_{t-1} and c_{t-2}. We eliminate W_t and W_{t-1} from (A.2) and obtain

$$c_t = \left\{(\beta R)^{1/A} + h\right\} c_{t-1} - h(\beta R)^{1/A} c_{t-2}, \tag{A.11}$$

with solution

$$c_t = \left\{\frac{c_0 - hc_{-1}}{(\beta R)^{1/A} - h}\right\}(\beta R)^{(t+1)/A} + \left\{\frac{(\beta R)^{1/A} c_{-1} - c_0}{(\beta R)^{1/A} - h}\right\} h^{t-1}. \tag{A.12}$$

Since $(\beta R)^{1/A} > 1$ and $0 \leq h < 1$, the first term in (A.12) dominates the second term. As $t \to \infty$, the ratio c_t/c_{t-1} tends to $(\beta R)^{1/A}$. It can also be shown that c_{t-1}/W_t tends to $(\beta R)^{-1/A} - R^{-1} > 0$ as $t \to \infty$.

The RRA coefficient is defined as

$$\begin{aligned} RRA &= \frac{-W_t V_{11}}{V_1} \\ &= \frac{A}{1 - hRc_{t-1}/(R-h)W_t} \quad \text{(by (A.9))} \\ &= \frac{A}{1 - h\left\{R(\beta R)^{-1/A} - 1\right\}/(R-h)} \quad \text{(in the steady state)}. \end{aligned} \tag{A.13}$$

We use the condition (A.5) to obtain

$$A \leq RRA \leq \frac{A}{1 - h(R-1)/(R-h)}. \tag{A.14}$$

The upper bound of the *RRA* coefficient is increasing in h. For example, for a large value of h, $h = 0.9$, and $R = 1.03$, we obtain $A \leq RRA \leq 1.3A$. We conclude that the *RRA* coefficient approximately equals the parameter A and that the approximation is not sensitive to the value of h.

Define the elasticity of consumption with the interest rate as

$$\begin{aligned} s &\equiv \frac{1}{c_t}\frac{\partial c_t}{\partial \ln R} \\ &= \left\{(\beta R)^{1/A}c_{t-1}/c_t - h(\beta R)^{1/A}c_{t-2}/c_t\right\}/A \\ &= \left\{1 - h(\beta R)^{-1/A}\right\}/A \quad \text{(in the steady state)}. \end{aligned} \tag{A.15}$$

If βR is approximately one, for example, we obtain $s \approx (1-h)/A$. We conclude that the elasticity of consumption is highly sensitive to h.

The product $s \times RRA$ in the steady state equals

$$s \times RRA = 1 - hR^{-1}. \tag{A.16}$$

Whereas the product equals one in the absence of habit persistence ($h = 0$), habit persistence drives a wedge between the *RRA* coefficient and the inverse of the elasticity of consumption in the sense that the product may be substantially below one.

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