

Time nonseparability in aggregate consumption

International evidence*

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'She ne'er had known pomp. Though't be temporal. Yet, if that quarrel, Fortune, do divorce
if from the bearer, 'tis a sufferance panging as soul and body's severing.'

William Shakespeare, from *Henry the VIII*: 2.3, 13–16.

We study consumption-based asset pricing models which allow for both habit persistence and durability of consumption goods, using quarterly consumption and asset return data for six countries. We estimate the parameters representing habit persistence or durability, risk aversion and time preference for each of the countries. We find that time-nonseparable preferences improve the fit of the model. When the nonseparability parameter is statistically significant, its magnitude indicates that the effect of habit persistence dominates the effect of durability in consumption expenditures. However, the international evidence for habit persistence is weaker than the evidence for the United States.

1. Introduction

The consumption-based asset pricing model of Lucas (1978) and Breeden (1979) has attracted extensive empirical study using data from the U.S. and

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abroad. The most common version of the model uses a time- and state-separable power utility function of a nondurable consumption good. The model has met with little empirical success. Its failures include the 'equity premium puzzle' [Mehra and Prescott (1985)] and rejections of the Euler equations [e.g. Hansen and Singleton (1982) and Hansen and Jagannathan (1991)].

Dunn and Singleton (1986) and Eichenbaum et al. (1988), among others, incorporated into their tests the idea that measured consumption expenditures represent the purchase of durable goods, which produce a flow of services over time. Durability introduces a form of nonseparability over time since the flow of services depends on past consumption expenditures. Using monthly U.S. data these studies found mixed evidence for durability. Durability improves the fit of the model, but in some cases the relevant durability coefficients are estimated to be of the wrong sign. Durability in consumption goods leads to a higher volatility of expenditures relative to the flow of services and the implied marginal utility of consumption. In the equity premium puzzle, Mehra and Prescott (1985) found that the volatility of the marginal utility is too low in a time-separable model. Therefore, durability has the wrong implication for the volatility of marginal utility.

Time-nonseparable utility exhibiting habit persistence was studied theoretically in Ryder and Heal (1973), Sundaresan (1989), Constantinides (1990), Novales (1990), and Detemple and Zapatero (1991), among others. If habit is formed over past consumption expenditures as in Constantinides (1990), the past expenditures enter the utility with a negative coefficient. Habit persistence tends to reduce the volatility of expenditures relative to the implied volatility of marginal utility. With habit persistence a consumer smooths consumption more than in a time-separable model. Ferson and Constantinides (1991) modeled both the durability of consumption expenditures and habit persistence, and empirically examined the Euler restrictions on U.S. monthly, quarterly and annual data. They showed that habit persistence and durability combine as opposing effects. Habit persistence tends to make the coefficients on lagged consumption expenditures negative while durability tends to reverse their sign. Ferson and Constantinides found that the evidence of durability in monthly U.S. data is weak, and that habit persistence dominates durability in quarterly and annual data. Ferson and Harvey (1992) examined a seasonal, non time-separable model and found evidence for seasonal habit persistence in U.S. quarterly data. Winder and Palm (1990) studied a linearized non time-separable model and found evidence for habit persistence in quarterly consumption data for the Netherlands.

Using quarterly data for Canada, France, Japan, the United Kingdom, the United States and Germany, we extend the empirical analysis of time-nonseparability in consumption-based asset pricing models. We allow for

both habit persistence and durability, estimating the parameters and testing the model separately for each of the countries. We find that abandoning time-separable preferences in favor of non time-separable preferences improves the fit of the model, as measured by the usual goodness-of-fit statistics. In those cases where the nonseparability parameter is statistically significant, it is always negative. This provides evidence that the effects of habit persistence dominate the effects of durability in consumption expenditures. However, the evidence for habit persistence is weaker in many countries than in the U.S.

In modelling a representative consumer economy with non time-separable preferences, we suppress a number of realistic and potentially empirically important considerations. These include transaction costs, constraints on consumer borrowing, and the inability to hedge consumer-specific endowment shocks due to incomplete markets. We suppress these aspects of the economy in order to assess the empirical relevance of non time-separable preferences.¹

The paper is organized as follows. The model is stated in section 2, the methodology is reviewed in section 3 and the data are described in section 4. The main empirical results are presented in section 5 and section 6 concludes the paper.

2. The model

Consider a stylized single-good economy in discrete time. Expenditures on the good at time t by a representative consumer are denoted by c_t . The good may be durable; specifically, the expenditures c_t at time t produce consumption services $\delta_\tau c_t$ at time $t+\tau$, where $\tau \geq 0$, $0 \leq \delta_\tau < 1$ and the infinite summation of the δ_τ equals one. The total flow of consumption services at time t is given by

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

A representative consumer's utility is defined over the flow of services, c_t^F . Habit persistence is modeled with a time-nonseparable von Neumann–Morgenstern utility function:

¹Transaction costs are examined by Cochrane and Hansen (1992), He and Modest (1991), Heaton and Lucas (1991), Luttmer (1991) and Scheinkman and Weiss (1986). Constraints on borrowing against future income are examined by Cochrane and Hansen (1992), He and Modest (1991) and Zeldes (1989). Methods for testing inequality restrictions, which could potentially be applied to liquidity constraints, are examined by Boudoukh et al. (1992). The inability to hedge against future endowment shocks is examined by Aiyagari and Gertler (1991), Bewley (1982), Constantinides and Duffie (1992), Imrihoroglu (1989), Lucas (1991), Mankiw (1986), Scheinkman (1989) and Telmer (1991).

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

The time-preference parameter is β . The habit parameter h , $0 \leq h < 1$, represents the fraction of the weighted sum of lagged consumption flows which establishes the subsistence level. The term $h \sum_{s=1}^{\infty} a_s c_{t-s}^F$ is the subsistence level. If $h=0$ there is no habit persistence and utility is time-separable in consumption flows (but not in consumption expenditures, unless δ is zero also). The parameters a_s , $0 \leq a_s < 1$, measure the persistence of lagged consumption flows in the subsistence level. The concavity parameter A , $A \geq 0$, represents the relative risk aversion coefficient (RRA) in the special case $h=0$. Constantinides (1990) argues that with habit persistence ($h>0$) the parameter A approximately equals the RRA coefficient.

Combining eqs. (1) and (2), the utility function may be written as

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

where

$$C_t = \delta_0 \sum_{\tau=0}^{\infty} b_{\tau} c_{t-\tau} \quad (4)$$

and

$$b_0 = 1, \quad b_{\tau} = \left(\delta_{\tau} - h \sum_{i=1}^{\tau} a_i \delta_{\tau-i} \right) / \delta_0, \quad \tau \geq 1. \quad (5)$$

A special case of the model illustrates the opposing forces of habit persistence and durability. Consider the case where depreciation of the good is exponential and habit persistence is exponential; that is, $\delta_{\tau} = (1-\delta)\delta^{\tau}$ and $a_s = (1-\alpha)\alpha^{s-1}$, where both δ and α are between 0 and 1. In this case

$$b_{\tau} = \delta^{\tau} \left[1 - h \left(\frac{1-\alpha}{\delta-\alpha} \right) \right] + \alpha^{\tau} \left[h \left(\frac{1-\alpha}{\delta-\alpha} \right) \right]. \quad (6)$$

If there is habit persistence but goods are not durable ($\delta=0$), then eq. (6) implies that $b_{\tau} = -(1-\alpha)h\alpha^{\tau-1} < 0$, $\tau \geq 1$. With no habit persistence ($h=0$) but with durable goods, $b_{\tau} = \delta^{\tau} > 0$, $\tau \geq 1$. When both effects are present, the coefficient b_{τ} is positive or negative depending on the relative magnitude of the durability parameter δ and the habit persistence parameter h and α . There are three possible cases. If $\delta \geq \alpha + h(1-\alpha)$, the coefficient b_{τ} is positive for all lags; if $\delta \leq h(1-\alpha)$ then b_{τ} is negative for all lags; finally, if

$h(1-\alpha) < \delta < \alpha + h(1-\alpha)$, b_τ is positive for recent lags and negative for distant ones. In other words, if habit persistence dominates durability for a given lag j ($b_j < 0$), then habit persistence must dominate durability at all greater lags ($b_k < 0$ for all $k > j$).

The intertemporal Euler equation follows from eq. (3) using a standard perturbation argument. Consider a reduction of the representative consumer's expenditures in period t from c_t to $c_t - \varepsilon$, $|\varepsilon| \ll 1$, the investment of ε in an asset with (stochastic) return R_{t+1} over one period and an increase of the consumption expenditures in period $t+1$ from c_{t+1} to $c_{t+1} + \varepsilon R_{t+1}$. The 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 eq. (4). Optimality of the consumption and investment plan requires that the expectation in period t of the utility of the consumption flows is maximized at $\varepsilon = 0$, leading to the Euler equation

$$E_t \left[\sum_{\tau=1}^{\infty} \beta^\tau (C_{t+\tau}/C_t)^{-A} (b_{\tau-1} R_{t+1} - b_\tau) - 1 \right] = 0, \quad (7)$$

where C_t is defined by eq. (4).

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

$$E_t [\beta (c_{t+1}/c_t)^{-A} R_{t+1} - 1] = 0. \quad (8)$$

Ferson and Constantinides (1991) were unable to estimate more than one nonseparable lag parameter, b_τ , with any precision and focused on a one-lag model of the Euler equation. That is, they assumed $b_\tau=0$ for all $\tau > 1$. Since consumption expenditure levels are highly correlated, it is not surprising that it is difficult to distinguish empirically which of several lags determines the subsistence level or to uncover a lag structure on several coefficients. This means that it is not possible to precisely estimate the half-lives of habit persistence and durability. Hansen and Jagannathan (1991), Eichenbaum and Hansen (1990), Gallant et al. (1990), and Ferson and Harvey (1992) also limited their attention to one-lag models. The one-lag model implies the following form of the Euler equation:

$$E_t \left(\beta \left[\left\{ \frac{c_{t+1} + b_1 c_t}{c_t + b_1 c_{t-1}} \right\}^{-A} + b_1 \beta \left\{ \frac{c_{t+2} + b_1 c_{t+1}}{c_t + b_1 c_{t-1}} \right\}^{-A} \right] R_{t+1} - b_1 \beta \left\{ \frac{c_{t+1} + b_1 c_t}{c_t + b_1 c_{t-1}} \right\}^{-A} - 1 \right) = 0. \quad (9)$$

Two technical issues arise in the models with nonseparable preferences. For the representative utility to be well-defined, the term $(c_{t+1} + b_1 c_t)$ must be positive with probability one. We cannot determine the support of the distribution of consumption, but we check to see that the term is positive for all realizations in our sample. We find that the condition is satisfied for all of the parameter values that we estimate.² A second issue is that if the marginal utility of the representative agent is not positive, then equilibrium may not exist. To address this concern, define the scaled marginal utility as

$$m_t = c_t^A dU_t/dc_t = E_t\{\gamma_{t+1}\}, \quad (10)$$

where

$$\gamma_{t+1} = [1 + b_1(c_{t-1}/c_t)]^{-A} + b_1\beta(b_1 + (c_{t+1}/c_t))^{-A},$$

U_t is the lifetime utility at date t , and $E_t\{\cdot\}$ denotes the conditional expectation at date t . We would like to establish that m_t is positive for all t . However, the marginal utility at date t depends on the conditional expected value of a function of future consumption. Positive marginal utility does not require that the realizations of γ_{t+1} are positive; only the conditional expectation of γ_{t+1} has to be positive.

We examine some necessary conditions for positive marginal utility.³ First, the unconditional mean of the expression (10) must be positive. We calculate the sample mean of (10), evaluated at our estimates of the parameter values, and find it to be positive for all cases. A more stringent analysis defines $x_t = [1 + b_1(c_{t-1}/c_t)]^{-A}$, and $y_{t+1} = [b_1\beta(b_1 + (c_{t+1}/c_t))^{-A}]$, for given values of the model parameters. Using a regression to approximate the conditional mean, by regressing y_{t+1} on x_t , we obtain estimates of $E\{y_{t+1} | x_t\} = \alpha_0 + \alpha_1 x_t$ and $E\{\gamma_{t+1} | x_t\} = \alpha_0 + (1 + \alpha_1)x_t$. We calculate a standard error for the conditional mean, given x_t , and use this to compute a t -statistic for each conditional expectation in the time series of the estimates of $\alpha_0 + (1 + \alpha_1)x_t$. None of the t -statistics are significantly below zero; the smallest is -0.57 . We conclude from these exercises that there is not a problem with negative marginal utility of the representative agent in our models, when the models are evaluated at the parameter values that we estimate in our data.⁴

2.1. A seasonal, non time-separable model

Ferson and Harvey (1992) study a model which allows for seasonal

²This should not be surprising given that the term $(c_{t+1} + b_1 c_t)$ is in the denominator of eq. (9). If this term starts in a positive region the algorithm used in the estimation will not wish to allow it to approach or pass through zero.

³We are grateful to a referee and to the editor for comments which stimulated these experiments.

⁴The t -statistics do not account for the fact that the smallest value in the time series is chosen, so they overstate the significance if evaluated in the standard fashion, which means that our conclusion is conservative.

durability or habit persistence. They apply the model to not seasonally adjusted quarterly data for the U.S., and find the results encouraging. The essence of the model is to assume that the subsistence level (in the case of habit persistence) or the flow of services (in the case of durability) depends only on consumption expenditures in the same quarter of the previous year. The model is parsimonious in that only a single parameter is used, both to control seasonality in the consumption data, and to model nonseparability. The seasonal, non time-separable model implies an Euler equation which is a simple modification of eq. (9):

$$E_t \left(\beta \left[\left\{ \frac{c_{t+1} + b_1 c_{t-3}}{c_t + b_1 c_{t-4}} \right\}^{-A} + b_1 \beta^4 \left\{ \frac{c_{t+5} + b_1 c_{t+1}}{c_t + b_1 c_{t-4}} \right\}^{-A} \right] R_{t+1} - b_1 \beta^4 \left\{ \frac{c_{t+4} + b_1 c_t}{c_t + b_1 c_{t-4}} \right\}^{-A} - 1 \right) = 0. \tag{11}$$

3. Methodology

The model parameters $\{\beta, A, b_1\}$ are estimated and the Euler equations are tested using Hansen's (1982) Generalized Method of Moments (GMM). First consider eq. (9), which defines an error term u_{it} for each asset i , $i = 1, \dots, N$, such that $E_t[u_{i(t+1)}] = 0$, where $E_t[\cdot]$ 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 $E[u_{t+1} | z_t] = 0$ and therefore $E[u_{t+1} \otimes z_t] = 0$, 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 $N \times L$ orthogonality conditions. The GMM estimates are based on minimizing the quadratic form $g'Wg$ where g is the $N \times L$ vector $(1/T) \sum_t [u_{t+1} \otimes z_t]$ and W is the inverse of a consistent estimate of the covariance matrix of these orthogonality conditions. Hansen (1982) discussed the formation of the weighting matrix W and provided 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, $N \times L$, 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 difference between the number of orthogonality conditions and the number of parameters. The parameters are $\{\beta, A, b_1\}$.

If we choose $A=0$ in the Euler eq. (9), we obtain $u_{t+1} = -(1 + b_1\beta) + \beta R_{t+1}(1 + b_1\beta)$. If we also choose b_1 and β such that $(1 + b_1\beta) = 0$ we obtain a trivial solution to the Euler equation. Following Eichenbaum and Hansen

(1990) and Ferson and Constantinides (1991), we divide the orthogonality conditions by $(1 + b_1\beta)$ when we estimate eq. (9) in order to avoid trivial solutions.

A negative value of the concavity parameter, A , is economically implausible because it implies negative risk aversion. Also, a maximum of the utility function may not exist if the parameter A is negative. In estimating the Euler equations we therefore restrict the domain of the parameter A to nonnegative values.⁵

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 implies that $E[u_{t+s}|u_t] = 0$, $s > 0$, and we say that u_t follows an MA(0) process. The time-separable model implies the null hypothesis,

$$H_0: b_s = 0, \quad s > 0, \quad \text{with an MA(0) error term } u_t.$$

In the one-lag model of eq. (9), where $b_s = 0$, $s \geq 2$, 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 neither is u_t , and the model does not imply that $E[u_{t+1}|u_t] = 0$ but implies that $E[u_{t+s}|u_t] = 0$, $s \geq 2$; therefore, u_t follows an MA(1) process. The one-lag model therefore implies the hypothesis

$$H_a: b_s = 0, \quad s \geq 2 \quad \text{with MA(1) error term } u_t.$$

The covariance matrix is adjusted to account for the moving-average terms as described by Hansen (1982). Note that when b_1 is zero, the model implies that the autocorrelation of the error becomes zero, hence the null hypothesis H_0 .

In the seasonal non time-separable model, the error term follows an MA(4) process when b_1 is not zero. By analogy with the nonseasonal model, we divide the orthogonality conditions by $(1 + b_1\beta^4)$ when we estimate eq. (11) in order to avoid trivial solutions, and we restrict the domain of the parameter A to positive numbers.

We model the representative consumer's decisions at fixed quarterly intervals and measure asset returns and consumption over the same intervals. Consumption decisions may actually be made more frequently. If decisions are made within the observation interval and the measured consumption expenditures are the sum of the expenditures over the interval, then the consumption data are said to be time-aggregated. Formally modeling time-aggregation in the nonlinear Euler equation is difficult. Theoretical results adjusting for time-aggregation are only available in the literature, imposing a first-order approximation on the marginal utility. With such a first-order

⁵We do this by dividing the error term u_{t+1} by $(1 + e^{5-10A})$. This is permissible by the same argument which justifies our scaling factor to avoid trivial solutions. Note that at given parameter values, the objective function, $g'Wg$, is invariant to the scaling factor.

approximation it can be shown that one effect of time-aggregation is to increase by one the order of the MA process followed by u_t .⁶ In the nonlinear Euler equation the results of time-aggregation are more complex. Therefore, under time aggregation, the residuals may appear to behave like a higher-order MA process even if the nonseparability parameter is zero. Time-aggregation can also induce a spurious correlation between the error terms and the information set for time t . Because of time-aggregation, variables in the market's information set at t which were not in the market's information set at $t-1$ may not be valid instruments for the equation $E_t[u_{t+1}] = 0$. We assess the sensitivity of our results to these effects 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 GMM are justified from asymptotic distribution theory. There is a natural concern about the properties of these procedures in small samples. Tauchen (1986), Kocherlakota (1990), and Mao (1991) provided simulation evidence for the time-separable model. Tauchen found that the test statistics perform well with as few as 50 annual observations, although he found a slight tendency to reject the model too infrequently. Kocherlakota and Mao, using different parameter values, found cases where the model is rejected too often. These studies suggest that the coefficient estimates and their asymptotic standard errors can be unreliable in small samples.

Ferson and Foerster (1992) studied the finite-sample properties of the GMM in a regression context with nonlinear, cross-equation restrictions. They found that a two-stage GMM approach, as described in Hansen and Singleton (1982), tends to reject the model too often in large systems while an iterated GMM approach provides more accurate test statistics. We use an iterated GMM approach in this study.⁷ Although we report the coefficient estimates and their asymptotic standard errors, we stress that the reliability of these estimates 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 which depend on the point estimates of the coefficients.

⁶Heaton (1993) models time aggregation using monthly U.S. data and a first-order approximation of the Euler equation in this paper. See also Ermini (1992).

⁷We construct the weighting matrix W using the parameter estimates from the n th stage, and use this matrix to find parameters for the stage $n+1$ which minimize the quadratic form. The new parameters are used to update the weighting matrix. The iterations continue until a minimum value of the quadratic form is obtained. In some cases, when an unconstrained optimization using this procedure chooses negative values of the parameter A , we use a variation of this procedure. We perform a grid search over the parameters A , β and b_1 and evaluate W at the parameter values which produce the approximate minimum value of the objective, subject to $A > 0$. We then minimize the objective function again, holding the W matrix fixed, to obtain the final parameter estimates.

4. The data

Our data consist of quarterly consumption expenditures and the real returns of common stocks and short-term bonds for six countries from 1970–1988. The countries are Canada, France, Japan, the United Kingdom, the United States and Germany.

4.1. *Stock return data*

The international common stock return data are from Morgan Stanley Capital International. Month-end indices are available from December, 1969. These are based on common stocks from a list of 1,476 (as of December, 1989) firms. The market value of these companies equals approximately 60% of the aggregate market value of the stock exchanges of the 19 countries for which the returns data are available. The selected stocks are generally those with large market capitalization. Investment companies and foreign-domiciled companies are excluded to minimize double-counting.⁸ Within each index, the stocks are value-weighted.⁹

4.2. *Short-term bond data*

We obtained short-term bond data from the Organization for Economic Cooperation and Development's (OECD) *Financial Statistics Monthly*. We selected 90–91 day Treasury bill returns when they were available (U.S., U.K. and Canada). When such rates were not available, other short-term rates were substituted. The nominal returns are converted to real returns, as described below.

4.3. *Consumption data*

The consumption data for all countries are from OECD's *Quarterly National Income and Product Accounts* publication. We use quarterly real per capita consumption expenditures for the period 1970 through 1989. The data are seasonally adjusted, with the exception of Germany and Japan, for which the reported series exhibit strong seasonalities. We use these data as they are reported for the seasonal non time-separable model and we seasonally adjust the two series for the nonseasonal model, using the following procedure. We regress the logarithm of the consumption expenditure level on a time-trend and dummy variables indicating the quarters. We add to the residuals the

⁸However, because of differences in corporate cross-holdings across countries, the relative value weights of MSCI indices may still be biased by double counting. See French and Poterba (1991), Harvey (1991), and Fedenia et al. (1991) for discussions.

⁹Further details of the construction of these indexes are found in Morgan Stanley's 'Capital International Perspective', quarterly issues, pp. 3–6.

sample mean of the logarithm of the consumption measure, plus the coefficient on the time-trend multiplied by the (mean-centered) time index, and we exponentiate the sum. This procedure removes seasonal fluctuations in the form of deterministic (percentage) quarterly shifts in expenditures, while retaining the overall mean and trend in the original expenditure series.

Total real consumption expenditures are divided by each country's population to obtain the real per capita consumption series. The population numbers are from the International Monetary Fund's data base. The population data are available on an annual basis, mid-year estimates. The annual figures were interpolated geometrically to get quarterly population numbers for each country.¹⁰

4.4. *Deflators*

The nominal returns for both the short-term bonds and the common stocks are converted to real returns using local currency consumption deflators. Consumption deflators for all countries are from OECD's *Quarterly National Income and Product Accounts* publication. All deflators are indexed in 1985 local currency units.

4.5. *Predetermined instruments*

Previous studies of consumption-based models showed that the results are sensitive to the choice of instruments. The earliest studies [e.g. Hansen and Singleton (1982)] concentrated on lagged consumption and returns as their instruments. However, measurement errors and time-aggregation can result in spurious correlations with lagged consumption and returns, and this can bias the coefficient estimates and test statistics. Ferson and Constantinides (1991) focused on financial variables, different from the lagged values of the model variables, as their instruments. They argued that such instruments, which predict both returns and consumption growth, are robust to spurious correlation problems and should provide powerful tests of the Euler equation restrictions. We adopt a similar strategy in this paper, and we investigate the sensitivity of our results to the choice of instruments.

Previous studies indicate which predetermined variables are useful in predicting international common stock and bond returns. Harvey (1991) showed that dividend yields and short-term interest rates predict stock returns in many countries. Solnik (1992) found economically significant predictability by such instruments. We include both of these variables in our

¹⁰The last two years (1988, 1989) were extrapolated from the 1987 series using the average growth rates for the previous ten years.

list.¹¹ We also include a term spread for each country, measured as the difference between the yield to maturity of a long-term bond and that of a short-term bond.

In addition to these local country-specific instruments, we employ a related set of financial instruments drawn from the U.S. market. The rationale for doing this is that the U.S. financial data available to us may be more accurate than financial data from the OECD and Citibase for the other countries. Campbell and Hamao (1992), Harvey (1991) and Ferson and Harvey (1993) found that financial instruments for the U.S. often predict stock returns in other countries. We use a one-month Treasury bill rate as our measure of the short-term U.S. interest rate and calculate a term spread as the difference between a long-term U.S. government bond return and the one-month bill. Finally, we use the dividend yield on the Standard and Poors' 500 stock index.

5. Empirical results

Tables with summary statistics for the basic data are available by request. The statistics reveal some interesting patterns. The first-order autocorrelations of quarterly consumption growth in Canada and the U.S. are positive, while the autocorrelations for the other four countries are negative. Non-separability is related to the autocorrelation in consumption. Durability of consumption expenditures induces negative autocorrelation. For example, a consumer purchasing an automobile in one period is unlikely to purchase another one for several periods. Habit persistence induces positive autocorrelation since the consumer maximizes utility by smoothing consumption more than would be optimal with time-separable preferences. On the one hand, measurement error in the levels of consumption can induce spurious negative autocorrelation in consumption growth. On the other hand, some components of consumption expenditures are calculated by interpolation, which can induce positive autocorrelation. Such measurement errors are troublesome in evaluating nonseparable consumption models, and may lead to erroneous conclusions about habit persistence and durability. As such problems are expected to be more pronounced if the recent-lag consumption growth is used as an instrument, we avoid using the most recent lag of consumption as an instrument.

¹¹The dividend yields are from Morgan Stanley Capital International. We use a three month bill yield directly when it is available (Canada, the U.S. and the U.K.) as an instrument, and we convert the yield into a quarterly rate of return when we use it as an asset return. For France, we use the day loan rate against private bills as an instrument and the one month PIBOR rate as the short term nominal return (prior to 1970:Q4 the PIBOR rate is not available so we use the day loan rate). For Germany we use the three-month FIBOR rate for both the return and the instrument. For Japan we use a call money rate until 1977:Q2 and the Gensaki rate after that, both as an instrument and as the nominal return.

We also examine the correlations across the countries of the asset returns and the instruments. The stock and bond return correlations are generally low, illustrating the well-known view that there are significant gains from international diversification. Also, the correlations among the per-capita consumption growth rates of the different countries are low. This further motivates our examination of the distinct Euler equations of a representative agent in each country, as it suggests that using an international representative agent's utility of consumption may not be useful for pricing the asset returns.

5.1. Diagnostic regressions

Expected returns and expected consumption growth rates were modeled as linear functions of predetermined instruments in early consumption-based asset pricing studies [e.g. Ferson (1983), Hansen and Singleton (1983)]. We report regressions of this type to further describe the data and help interpret the results, but our formal tests do not assume that expected returns or consumption growth rates are linear functions of these variables. Table 1 shows time-series regressions, using the U.S. instruments to predict the future variables. The first panel shows regressions for the stock returns, the second panel for the bond returns and the third panel for the consumption growth rates. The regressors in each case are the lagged nominal stock market return for the U.S., the U.S. one-month bill yield, lagged U.S. consumption growth, the U.S. term spread and the dividend yield of the Standard and Poors 500. To be conservative, all of the regressors are lagged twice in relation to the dependent variables. The results are representative of the patterns that we find using other choices for the lagged instrumental variables.

In the first panel of table 1 there is evidence that quarterly real stock returns in the six countries are predictable to some extent using the lagged instruments. The Treasury bill and dividend yield are the instruments with the most forecast power, which is consistent with earlier work [e.g. Harvey (1991), Solnik (1992)]. The S&P 500 dividend yield has a positive and significant coefficient for five of the six countries. The Treasury bill rate has a negative coefficient in all countries, but is significant only in the U.S. and Canada.

The second panel of table 1 shows the regressions for the short-term real bond returns. The term spread and the short-term bill rate are powerful predictors in these regressions, although the dividend yield and the lagged stock return also have *t*-ratios larger than two in some cases. The term spread has a negative coefficient in each regression. The lagged consumption growth also has negative coefficients in all regressions, which suggests that relatively high real returns on short term bonds are associated with recessions or times of recent poor economic growth. The high adjusted *R*-

Table 1
 Regressions of real returns and consumption growth on lagged variables.^a

Instrument:	a_i	CGROW	RSTOX	TBILL	DIVYLD	TERM	$R^{2\%}$	r_1
Real stock returns								
U.S.	-0.11 (2.30)	-0.15 (0.10)	-0.10 (0.94)	-1.18 (2.24)	5.66 (3.89)	0.52 (0.85)	17.6	0.13
U.K.	-0.14 (1.95)	1.27 (0.68)	-0.22 (0.73)	-1.78 (1.57)	8.08 (2.30)	0.17 (0.16)	11.2	0.07
Japan	0.12 (1.68)	1.42 (1.06)	-0.13 (0.90)	-0.20 (0.30)	-1.32 (3.66)	1.12 (2.02)	6.4	-0.06
Germany	-0.12 (1.99)	0.73 (0.70)	0.01 (0.11)	-0.34 (0.64)	3.66 (1.94)	1.13 (2.02)	6.4	-0.06
France	-0.14 (1.90)	2.14 (1.31)	-0.27 (1.93)	-1.05 (1.24)	6.28 (2.74)	0.57 (0.72)	13.5	-0.03
Canada	-0.10 (1.79)	1.50 (1.28)	-0.05 (0.52)	-1.72 (2.54)	6.66 (3.88)	-0.49 (0.64)	15.3	0.21
Real short-term bond returns								
U.S.	0.015 (4.95)	-0.07 (0.84)	0.0005 (0.08)	0.18 (5.53)	0.32 (3.81)	-0.30 (7.51)	83.8	0.25
U.K.	0.04 (4.25)	-0.53 (2.41)	-0.003 (0.12)	-0.18 (1.69)	0.97 (3.16)	-0.50 (5.07)	36.9	0.57
Japan	0.06 (5.68)	-0.50 (2.49)	-0.05 (2.83)	-0.15 (1.43)	-0.20 (0.54)	-0.46 (3.73)	27.4	0.52
Germany	0.04 (6.36)	-0.36 (2.77)	-0.03 (2.60)	0.23 (3.64)	-0.59 (3.05)	-0.19 (2.63)	40.2	0.27
France	0.02 (2.95)	-0.30 (2.11)	-0.01 (0.95)	0.27 (4.80)	0.29 (1.52)	-0.18 (2.34)	58.1	0.50
Canada	-0.003 (0.85)	-0.12 (1.52)	0.003 (0.30)	0.27 (5.20)	0.63 (5.00)	-0.10 (1.85)	78.3	0.49

	0.01 (1.44)	-0.02 (0.27)	0.02 (2.17)	-0.05 (0.97)	-0.05 (0.33)	0.20 (4.13)	37.7	-0.06
U.S.								
U.K.	0.01 (1.57)	0.10 (0.57)	-0.01 (0.36)	-0.03 (0.41)	-0.14 (0.51)	0.12 (1.11)	0.0	-0.20
Japan	0.01 (1.02)	0.07 (0.47)	-0.04 (2.77)	-0.38 (0.51)	-0.26 (0.90)	0.03 (0.30)	1.8	-0.22
Germany	0.01 (1.38)	-0.14 (0.70)	-0.01 (0.42)	-0.14 (1.65)	-0.03 (0.13)	0.09 (0.90)	5.0	-0.40
France	0.01 (2.28)	-0.08 (0.82)	0.01 (0.99)	-0.03 (0.60)	-0.15 (0.76)	0.02 (0.45)	2.7	-0.23
Canada	0.03 (6.02)	-0.11 (0.98)	0.02 (1.18)	-0.07 (1.01)	-0.45 (2.38)	-0.00 (0.02)	26.6	-0.21

^aQuarterly data from 1970:Q4-1988:Q4 (72 observations) are used (for Japan, the data are for 1970:Q4-1988:Q1 and there are 69 observations). The regression is

$$R_{i,t+1} = a_i + \beta_{i1}Z_{1,t-1} + \dots + \beta_{i5}Z_{5,t-1} + \varepsilon_{i,t+1}$$

where R_i is the real asset return measured in local currency units or, the real consumption growth and the Z_j are the lagged instruments. The instruments are (1) the real, lagged per capita growth of total personal consumption expenditures in the United States, denoted by *CGROW*; (2) the lagged nominal return of the U.S. stock market index, denoted by *RSTOX*; (3) the lagged nominal return of a U.S. Treasury bill, denoted *TBILL*; (4) the Divided yield of the Standard and Poors 500 stock index, denoted *DIVYLD*; and (5) a U.S. term spread, denoted *TERM*, and measured as the difference between the yield to maturity of a long term U.S. government bond and a one-month Treasury bill. All of the instruments are lagged two quarters relative to the dependent variable. The regression betas are listed in the table. The absolute values of the heteroskedasticity-consistent *t*-statistics are in parentheses. r_1 is the first order autocorrelation of the regression residual. The *R*-squares are adjusted for degrees of freedom.

squares for the short-term bond returns in table 1 are remarkable. This is a peculiarity of the 1970–1988 sample period, when interest rates were high, volatile and highly autocorrelated in many countries. We checked these regressions against alternative data for the U.S. over this period and found similar results. Although the short-term bond returns are highly predictable in the quarterly data over this period, they do not seem to contain a unit root.¹²

The bottom panel of table 1 reports regressions of the continuously-compounded consumption growth rates on the lagged instruments. The adjusted *R*-squares of the regressions are in excess of 25% for the U.S. and Canada, but are much lower for the other countries. While we do not use these growth rates directly in our tests of the models, the regressions provide information about the likely power of our tests.¹³

5.2. *Empirical tests using own-country instruments*

For each country the asset system corresponding to (9) consists of the local value-weighted common stock index return and the short-term bond return. Consumption is the local real, per-capita consumption. The instruments are six: a constant and five own-country variables. The five variables are the real consumption growth rate lagged twice and the nominal stock market return, short-term bond return, term spread and dividend yield, each lagged once. We use nominal, instead of deflated, financial instruments to avoid the possibility of spurious correlation introduced by deflating both the returns and the instruments by the same deflator series.¹⁴ The errors

¹²We examined the U.S. three-month spot rates from CRSP as an alternative to the OECD data. The series plots tracked each other closely. Regressing the real returns using this alternative nominal rate series on the instruments in table 3 produced an adjusted *R*-square equal to 81%. We estimated an AR(3) model for the real returns. The first-order lagged coefficient in the model was 0.69, with a (heteroskedasticity-consistent) standard error equal to 0.12. The adjusted *R*-square of this model is 72.9%.

¹³The Euler equation implies that the product of a real asset return and the marginal rate of substitution of consumption should not be predictable by lagged instruments. In the time-separable model the marginal rate of substitution is one plus the growth rate of consumption raised to the power $-A$. In the nonseparable model the marginal rate of substitution is a more complicated function of consumption. If the instruments can predict consumption growth and returns, the tests should have some power.

¹⁴We use consumption at lag 2 only while allowing the financial instruments to enter at lag 1 in an attempt to avoid spurious correlations while obtaining good predictive power. Spurious correlation due to time aggregation is a problem for the consumption data, because time aggregation implies that the same underlying 'true' consumption appears in the time-averaged consumption data over two adjacent periods. Time aggregation also can induce spurious correlation with other variables at lag one, to the extent that the underlying true contemporaneous consumption in the future time-averaged consumption data is correlated with the financial variables. As this correlation is likely to be much smaller than 1.0, the spurious correlation induced by time aggregation is likely to be less of a problem for the lagged financial instruments than for lagged consumption expenditures.

u_{t+1} in the Euler equation $E_t[u_{t+1}] = 0$ are assumed to follow an MA(0) process in the time-separable model, an MA(1) process in the nonseparable model, and an MA(4) process in the seasonal non time-separable model.

The results are presented in table 2. The notation $b_1 \equiv 0$ means that we set the nonseparability parameter equal to zero, thereby estimating and testing the time-separable model. For the time-separable model the subjective discount rate (β) is estimated precisely and is less than 1.0 in all countries. The point estimates of the concavity parameter, subject to the restriction that $A \geq 0$, are all around one or less. The right tail p -values for the goodness-of-fit tests are less than 5% in the U.K. and Japan; they are between 5% and 10% in the U.S.; and they exceed 10% in France, Canada, and West Germany.

In the second row of each panel of table 2 the nonseparability parameter (b_1) is estimated along with β and A , in the nonseasonal, non time-separable model (9). The model is not rejected by the goodness-of-fit test in any of the six countries, the p -values exceeding 15%. The point estimates of A are plausible but the standard errors are large. The estimates of the nonseparability parameter are all negative, but significant only in the U.S. The evidence for habit persistence in the U.S. is consistent with the results of Ferson and Constantinides (1991).

The results for the seasonal non time-separable model are reported for Germany and Japan in the third row of their panels in table 2. The model is not rejected by the goodness-of-fit tests and the point estimates of the nonseparability parameter are negative. The standard errors of the coefficients are typically much smaller than in the nonseasonal models, and the nonseparability parameters are many standard errors away from zero. These results are similar to the findings of Ferson and Harvey (1992) for U.S. data.

Fig. 1 illustrates the sensitivity of the model to the value of the nonseparability parameter. The values of the objective function are illustrated when the objective is minimized over the choice of A and β , for given values of b_1 . An MA(1) weighting matrix [from Hansen (1982)] is used in the nonseasonal models and an MA(4) is used in the seasonal model. The objective function is highly nonlinear in the parameter b_1 .¹⁵ Often there are local minima both in the region of durability ($b_1 > 0$) and in the region of habit persistence ($b_1 < 0$).

Ferson and Constantinides (1991) found local minima in the regions of durability and habit persistence, using U.S. monthly, quarterly and annual data. Heaton (1993) found evidence in U.S. data which he interpreted as consistent with the presence of durability, which operates with a relatively short half life, and also of habit persistence which operates with a longer half

¹⁵Recall that the objective function is a quadratic form in the Euler equation errors, which are themselves nonlinear functions of the model parameters.

Table 2

Test results for time separable and time nonseparable models using own-country instrumental variables.^a

Country	β	A	b_1	χ^2	p -value
United States	0.979 (0.011)	1.10 (1.58)	$\equiv 0^b$	16.79	0.079
	0.754 (0.240)	3.11 (21.4)	-0.92 (0.40)	13.24	0.152
United Kingdom	0.967 (0.007)	0.87 (0.89)	$\equiv 0^b$	21.82	0.016
	0.547 (0.741)	6.30 (124.5)	-0.82 (3.08)	1.07	0.999
France	0.962 (0.020)	1.07 (3.45)	$\equiv 0^b$	9.03	0.529
	0.94 (0.27)	1.18 (52.4)	-0.69 (5.67)	3.25	0.954
Canada	0.978 (0.012)	0.98 (1.19)	$\equiv 0^b$	13.77	0.184
	0.16 (7.14)	12.76 (1,018.0)	-0.93 (5.04)	0.01	0.999
Germany	0.983 (0.017)	1.17 (4.44)	$\equiv 0^b$	8.36	0.594
	0.900 (1.15)	2.09 (198.1)	-0.63 (20.4)	0.09	0.999
Seasonal model ^d	0.990 (0.021)	1.08 (1.58)	-0.74 (0.19)	7.54	0.581
Japan ^c	0.970 (0.001)	0.35 (0.40)	$\equiv 0^b$	21.70	0.017
	0.918 (0.036)	3.10 (7.39)	-0.64 (0.42)	10.98	0.277
Seasonal model ^e	0.983 (0.007)	0.76 (0.26)	-0.57 (0.08)	10.35	0.323

^aModels with habit persistence or durability of consumption expenditures using quarterly returns and consumption data for individual countries. The returns are measured for 1970:Q.4–1988:Q.4 (72 observations). A is the concavity parameter, β is the subjective discount rate 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. The asset returns are a value-weighted common stock return index and a short term bond return for each country. Both are measured in local currency units and are deflated by the consumption price deflator for the country. The instruments include a constant and the consumption growth measure at lag two. The instruments also include the local nominal stock return, the local nominal bond return, the local dividend yield and the local term spread, all measured at lag one.

^bAn ' $\equiv 0$ ' indicates that the parameter is set to zero.

^cThe non-seasonal model for Japan uses return data for 1970:Q4–1988:Q1 (69 observations).

^dThe seasonal model for Germany uses return data for 1971:Q3–1988:Q2 (68 observations).

^eThe seasonal model for Japan uses return data for 1971:Q3–1987:Q2 (64 observations).

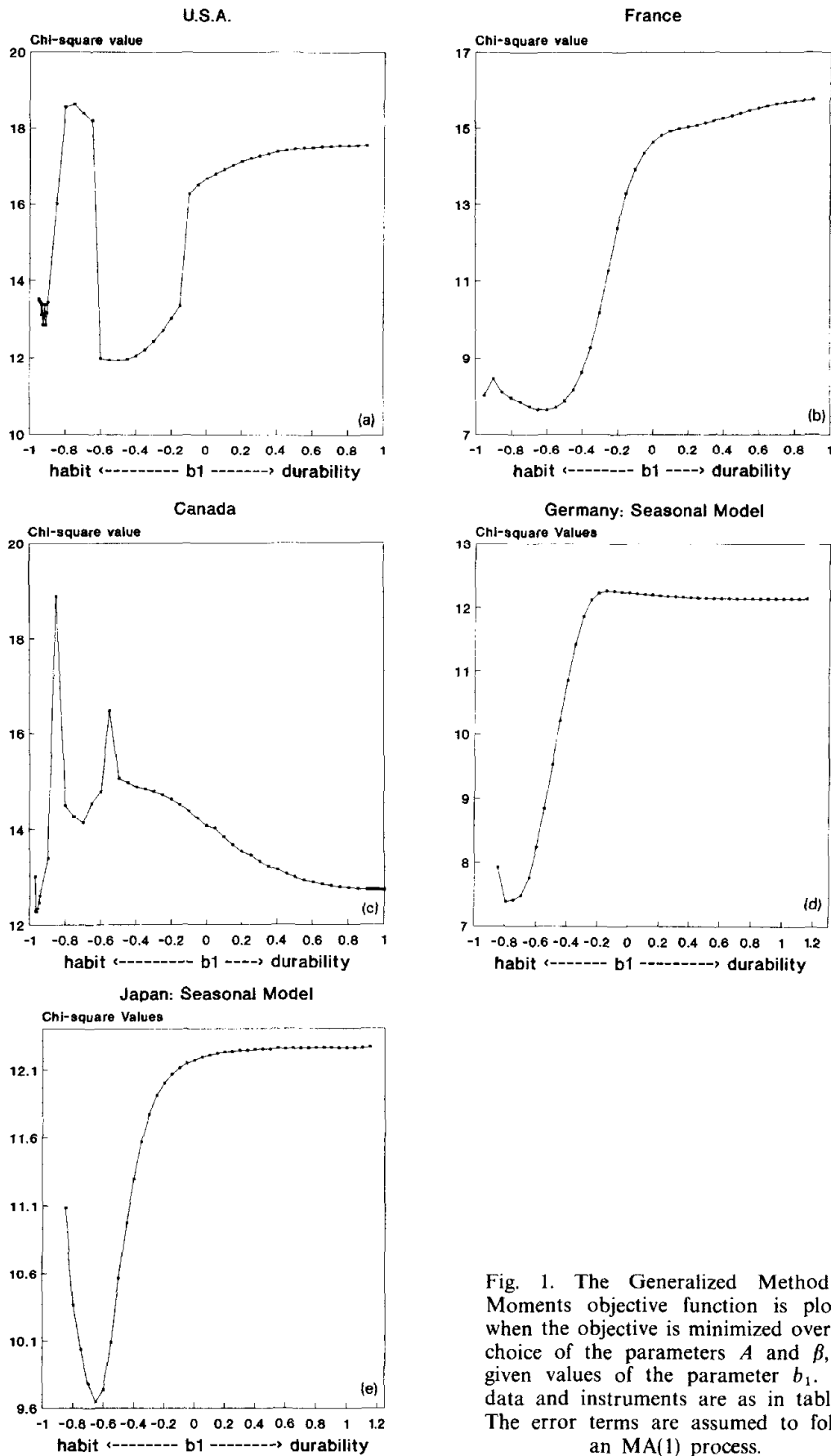


Fig. 1. The Generalized Method of Moments objective function is plotted when the objective is minimized over the choice of the parameters A and β , for given values of the parameter b_1 . The data and instruments are as in table 2. The error terms are assumed to follow an MA(1) process.

life. Constantinides (1990) suggested that habit persistence in the U.S. may have a half-life in excess of one year. The presence of local minima may suggest that both the effects of durability and of habit persistence are present in the data for the other countries as well. Since our model uses only the single lag coefficient b_1 , it can not discriminate between the two effects when both are present.

Time aggregation implies that the first lagged values of the instruments, as we used them in table 2, may not be valid instruments. Time aggregation may also induce a higher order moving average process in the error terms. Therefore, we modify the procedure by imposing an MA(1) process on the error terms in the time-separable model and an MA(2) process in the nonseasonal nonseparable model. (These results are available by request.) Neither the time-separable nor the nonseparable model is rejected in any of the countries. The coefficient b_1 is significantly negative in Japan and in the U.S., but is insignificantly different from zero in the other countries. The point estimates of β and the nonseparability parameter, b_1 , are fairly stable across these experiments, but the estimates of the risk aversion parameter, and especially the standard errors of the coefficients, are not. The larger standard errors may in this case be attributed to the higher order moving average terms.

In a second experiment we lag all of the instruments twice (recall that in table 2 we lagged consumption growth twice and the financial variables only once. We use an MA(0) process on the error terms in the time-separable model and an MA(1) process in the nonseparable model (tables of these results are available by request). We find that none of the models are rejected for any of the countries by the goodness-of-fit tests. The estimates of the coefficient b_1 are negative in all countries but significant only for Japan.

5.3. Tests using U.S. instruments

The adjusted R -squares reported in table 1 suggest that U.S. instruments have reasonable forecasting power for the return and consumption growth of other countries. This motivates a replication of the experiments, replacing the own-country instruments with the U.S. instruments. The results are reported in table 3.

Using the U.S. instruments, the time-separable model is rejected by the goodness-of-fit tests only in Japan, where the right-tail p -value is 0.041. The non time-separable model is not rejected at the 5% level, the p -value for Japan being 0.092. However, the standard errors of the nonseparability parameter are very large in the nonseasonal, non time-separable model.¹⁶

¹⁶Using the U.S. instruments at lag two, we find that the time-separable model is rejected by the goodness-of-fit tests at the 5% level in the U.K. and West Germany, and at the 10% level also in the U.S. and Japan. The nonseparable model is not rejected in any country at the 10%

The seasonal non time-separable model is not rejected, the p -values being 0.48 or higher. While the standard errors of the risk aversion parameter are large in this model, the nonseparability parameter seems to be precisely estimated, at -0.7 for both Germany and Japan.

While the point estimates of b_1 tend to be fairly stable across the experiments, the estimates of the other parameters and the standard errors tend to be unstable. The instability of the results across the specifications may not be surprising, given a highly nonlinear model and fairly small sample sizes. Such instability may be further interpreted as indicating that the model is misspecified. The standard errors are often quite large in the non time-separable models, indicating that there is not enough information in these data to reliably infer the values of all parameters simultaneously.

6. Conclusions

We estimated and tested consumption-based asset pricing models which incorporate both durability of goods and habit persistence in the preferences of country-specific representative consumers. By postulating a representative consumer in each country, as opposed to a global representative consumer, we recognized the incompleteness of markets across countries, but implicitly assumed that markets are effectively complete within each country. We broadened the investigation of Ferson and Constantinides (1991) in U.S. data to the representative consumers of six major countries. A series of controlled experiments was conducted to evaluate the sensitivity of the results to different instruments, timing conventions for the instruments and assumptions about the autocorrelation of the model error terms.

We found evidence in favor of habit persistence against a net effect of durability in consumption expenditures in the U.S., and to a lesser extent for Japan. The point estimates of the nonseparability parameter are in the direction of habit persistence, as opposed to durability, in all of the countries and in most of the experiments. Goodness-of-fit tests indicate improved fit when the models allow for habit persistence. However, the parameter estimates are imprecise and unstable across the experiments. The low power of the tests may be partly due to the poor quality of the available data on international consumption and returns. The instability of the results across experiments also suggests that the model is misspecified. Future research should attempt to incorporate more realistic considerations such as more general utility functions, market imperfections such as liquidity constraints and uninsurable risks.

level. The nonseparability parameter is significantly different from zero and negative in the U.S., Canada and Germany. Using a higher order moving average term none of the models are rejected by the goodness-of-fit tests, and the nonseparability parameter is significantly negative only in the U.S.

Table 3

Test results using own-country financial assets and U.S. instrumental variables.^a

Country	β	A	b_1	χ^2	p -value
United States	0.975 (0.010)	1.02 (1.50)	$\equiv 0^b$	14.64	0.146
	0.942 (0.032)	0.23 (2.02)	-0.94 (0.32)	10.65	0.301
United Kingdom	0.969 (0.021)	1.08 (2.75)	$\equiv 0$	9.09	0.523
	0.912 (0.349)	1.21 (52.1)	-0.72 (5.61)	3.83	0.922
France	0.959 (0.025)	1.01 (4.32)	$\equiv 0$	7.44	0.683
	0.964 (0.055)	1.08 (9.93)	-0.60 (1.52)	3.48	0.943
Canada	0.975 (0.012)	0.91 (1.28)	$\equiv 0$	10.64	0.386
	1.048 (0.976)	1.20 (123.1)	0.99 (1,002.4)	0.41	0.999
Germany	0.976 (0.016)	1.19 (3.37)	$\equiv 0$	5.16	0.880
	0.984 (0.023)	1.18 (8.43)	-0.57 (1.27)	3.17	0.957
Seasonal model ^d	1.004 (0.020)	1.04 (1.38)	-0.72 (0.16)	8.46	0.488
Japan ^c	0.970 (0.002)	0.65 (0.44)	$\equiv 0$	18.94	0.041
	0.972 (0.005)	0.09 (1.38)	-0.75 (1.70)	14.95	0.092
Seasonal model ^e	0.986 (0.030)	1.31 (1.65)	-0.72 (0.19)	7.77	0.558

^aModels with habit persistence or durability of consumption expenditures using quarterly returns and consumption data for individual countries. The returns are measured for 1970:Q4–1988:Q4 (72 observations). The model and the data are the same as in table 2, except that the instruments include a constant and the consumption growth measure for the U.S. measured at lag two. Also included are the U.S. nominal stock return, the U.S. 1-month Treasury bill rate, the dividend yield of the Standard and Poors 500 stock index, and the U.S. terms spread (U.S. long term bond yield less 1 month bill rate), all measured at lag one.

^bAn ' $\equiv 0$ ' indicates that the parameter is set to zero.

^cThe non-seasonal model for Japan uses return data for 1970:Q4–1988:Q1 (69 observations).

^dThe seasonal model for Germany uses return data for 1971:Q3–1988:Q2 (68 observations).

^eThe seasonal model for Japan uses return data for 1971:Q3–1987:Q2 (64 observations).

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