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An empirical investigation of asset pricing models using Japanese stock market data

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This article examines the empirical performance of the following asset pricing models: (i) a time-separable model; (ii) Abel's model with a consumption externality; (iii) the time non-separable model; (iv) the consumption-based recursive utility model; and (v) the ad-hoc factor pricing model. The testing framework includes the Hansen–Jagannathan volatility bounds test, the Hansen–Jagannathan specification error test and Euler equation-based generalized method of moments estimation. The test evidence indicates that habit forming preferences are empirically supportable and provide a good characterization of the Japanese security market data. (JEL G12).

The time and state-separable isoelastic utility asset pricing model has failed to account for one important feature in the US security markets: consumption growth varies too little to explain the observed equity premium unless relative risk aversion is high. One possible explanation for this failure is that the maintained specification of preferences is too rigid. While departures from the standard model have been studied using US data, this paper investigates the empirical performance of intertemporal asset pricing models using Japanese stock market data. In particular, we study and contrast the empirical fit of the standard expected utility model relative to: (i) Abel's (1990) model that allows for consumption externality; (ii) the time non-separable model as in Ferson and Constantinides (1991); (iii) the recursive utility model as in Epstein and

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Zin (1991); and (iv) the ad-hoc factor pricing model. Our testing methods include the Hansen and Jagannathan (1991) volatility bounds tests, the Hansen and Jagannathan (1994) specification error tests, and generalized methods of moment (GMM) tests of the asset pricing restrictions. The purpose of this study is to provide an understanding of the dimensions along which existing asset-pricing models may be failing in Japan and to contrast the results to what we know about US security markets. Given the recent academic interest in the Japanese capital markets, the preceding empirical exercises can guide the search for better performing intertemporal asset pricing models.

In the existing literature, GMM-based tests of the Euler equations form the sole basis for the empirical comparison of nested asset pricing models. In their influential study, Ferson and Constantinides (1991) demonstrate that the time non-separable model performs better empirically than the time-separable model. In particular the asset pricing model with non-separabilities in consumption has incremental power in explaining the mean and standard deviation of the equity premium (see also, Constantinides, 1990). However, when models are non-nested, the GMM methodology offers little guidance on the relative usefulness of asset pricing models. As the simulation-based studies of Ferson and Foerster (1994) and Hansen *et al.* (1994) emphasize, the GMM estimator possesses undesirable finite sample properties and the GMM criterion function rejects *too often*. The exclusive reliance on GMM estimation can therefore lead to misleading inferences with regard to the validity of asset pricing models.

The recent work of Hansen and Jagannathan (1994) and Hansen *et al.* (1995) provides a comprehensive framework whereby competing asset pricing models can be compared and contrasted. In their empirical methodology, it is understood that a stochastic discount factor will not necessarily explain 100% of the variation in security prices. They derive a specification test that calculates the *maximum pricing error* associated with a stochastic discount factor. This specification error test calculates the least square distance between the candidate stochastic discount factor under study and the set of admissible stochastic discount factors. The stochastic discount factor with the least distance measure performs better in the sample. Since the parameters determining the stochastic discount factor can be estimated by minimizing the specification error, the specification error test yields crucial comparisons of the preference parameters with the GMM counterpart.

For any class of intertemporal asset pricing models, the general treatment in Hansen and Jagannathan (1991) permits the investigator to restrict the admissible region for the mean and standard deviation of the stochastic discount factors. If the stochastic discount factor (corresponding to any candidate asset pricing model) does not generate enough volatility, it will lie outside the Hansen–Jagannathan *volatility bound* and the pricing model can be declared inconsistent with the security market data. Thus, the Euler asset pricing restriction implied by each asset pricing model together with the unconditional moments of the Japanese security market data, implies a lower bound on the volatility of the stochastic discount factor. Since the volatility bounds can be constructed directly from the mean–standard deviation frontier for asset returns, the focal point of this component of the empirical investigation is the 'market price of risk'. Hansen and Jagannathan (1991) apply this approach to assess the plausibility of the structural parameters that are consistent with the observed dynamics of asset prices. Following their pioneering work, the computation of the lower volatility bound is now a standard practice for the evaluation of asset pricing models (e.g. Ferson, 1994). Application of their volatility bounds diagnostic to the Japanese security markets indicates that the stochastic discount factor implied from the security markets is extremely volatile. Bakshi and Chen (1996) and Cochrane and Hansen (1992) nonetheless point out that the inability to generate volatile stochastic discount factors is a limitation of most existing asset pricing models.

The emphasis on the Japanese security markets is motivated by two principal considerations: (a) the Japanese stock market is second in size only to the United States; and (b) the large fluctuations in the Japanese stock market should pose an even greater challenge for any reasonably parameterized asset pricing model. Formal tests conducted by minimizing the specification error allow us to conclude that the time non-separable model and the ad hoc factor pricing model (with rate of return on the market portfolio, consumption growth, and lagged consumption growth as factors) offer the best empirical fit to the Japanese stock and bond market data. Overall the parameter estimates are quite supportive of the habit formation model. Even though the elasticity of intertemporal substitution is no longer constrained to be the reciprocal of the risk aversion parameter, the performance of the recursive utility asset pricing model is relatively unimpressive. For instance, its specification error differs only slightly, relative to the time separable asset pricing model, and is much larger than the ad hoc factor and the time non-separable asset pricing models.

Three main results emerge from the Hansen–Jagannathan (1991) volatility bounds diagnostic: (i) The minimum volatility of the bound is 25% and corresponds to a mean discount factor of 0.9975. (ii) For an economically plausible coefficient of relative risk aversion, only a moderate amount of habit formation is consistent with monthly Japanese consumption data. (iii) With the time-separable asset pricing model, a large risk aversion parameter (in excess of 12) is required to satisfy the Hansen–Jagannathan bounds. The results from the GMM estimation based upon monthly data support *neither* durability nor habit formation. This conclusion is robust to the set of conditioning variables. With quarterly consumption expenditures data, habit formation dominates the effect of durability. The goodness-of-fit Hansen's *J*-test (1982) is rejected for most asset pricing models and the parameter estimates are often imprecise.

The paper proceeds as follows. Section I outlines the preference structure and the testable restrictions for a broad class of asset pricing models. Section II first presents the Hansen and Jagannathan (1991) volatility bound diagnostic and the statistical test of the volatility bound proposed by Cecchetti *et al.* (1994). The specification error test of Hansen and Jagannathan (1994) is then implemented. This exercise permits us to determine the relative usefulness of non-nested preference structures. In Section III, the over-identifying restrictions implied by each asset pricing model are tested via Hansen's (1982) GMM.

Concluding remarks are offered in Section IV. An Appendix contains a description of the data.

I. Preferences and asset pricing

Most asset pricing models can be expressed in terms of their *stochastic discount factor* or intertemporal marginal rates of substitution (IMRS). Ross (1978), Hansen and Richard (1987), and Hansen and Jagannathan (1991) have demonstrated that when the *Law of One Price* holds:

$$\langle 1 \rangle \qquad E_t(m_{t+1}R_{j,t+1}) = 1 \qquad \forall t.$$

Here $R_{j,t+1}$ is the gross real return on asset j and E_t denotes expectation with respect to the conditioning information. Asset pricing models propose different stochastic discount factors (denoted by m_{t+1}) but all of those models must satisfy ⟨1⟩. The stochastic discount factors from the following models will prove useful for the empirical work to follow.

I.A. Model 1: Time non-separable utility

Ferson and Constantinides (1991), and Hansen and Jagannathan (1991) consider a model where time-t utility depends not only on current consumption (c_t) but also on its most recent past.[1] The stochastic discount factor for this model is

$$\langle 2 \rangle \qquad m_{t+1} = \beta \frac{(c_{t+1} - bc_t)^{-\gamma} - b\beta E_{t+1}(c_{t+2} - bc_{t+1})^{-\gamma}}{(c_t - bc_{t-1})^{-\gamma} - b\beta E_t(c_{t+1} - bc_t)^{-\gamma}},$$

where β represents the subjective discount factor. Ferson and Constantinides (1991) and Heaton (1995) allow more complex lag structures in which past consumption can affect future consumption. We nonetheless choose to focus on the simple representation in ⟨2⟩ since the empirical implementation of the non-separable model with additional lags often leads to unreliable parameter estimates (see, for example, Ferson and Constantinides, 1991). The sign of the non-separable parameter, b, is a critical determinant of whether durability or habit formation dominates.

- If $b > 0$, then non-separability in consumption supports habit formation. An increase in the past consumption with no change in time-t consumption causes the period utility to decrease. Constantinides (1990) illustrated how γ closely approximates the coefficient of relative risk aversion (RRA), and found that habit persistence is useful in explaining both the mean and standard deviation of the equity premium.
- Otherwise, if $b < 0$, past consumption provides utility in the current period and consumption is durable (e.g. Dunn and Singleton, 1986; Eichenbaum and Hansen, 1990).
- If $b = 0$, the non-separable model reduces to the standard *time-separable*

expected utility model. For this special case, the stochastic discount factor is

$$\langle 3\rangle \qquad m_{t+1} = \beta\left(\frac{c_{t+1}}{c_t}\right)^{-\gamma}.$$

I.B. Model 2: Consumption externality

In Abel (1990) the consumer–investor is not only concerned with his time-t consumption, but also with his consumption expenditures relative to the lagged cross-sectional average level of consumption.[2] When the consumption externality arises from lagged aggregate consumption, the stochastic discount factor takes the form

$$\langle 4\rangle \qquad m_{t+1} = \beta\left(\frac{c_{t+1}}{c_t}\right)^{-\gamma}\left(\frac{c_t}{c_{t-1}}\right)^{-(1-\gamma)},$$

where γ represents the concavity parameter, and the second term in parentheses captures the effect due to '*catching up with the Joneses*'.

I.C. Model 3: Recursive utility model

The third empirical model to be tested is a variant of the Kreps–Porteus and Epstein and Zin (1991) recursive utility models. The driving force in this model is the assumption that investors, in evaluating their consumption stream, care about the timing of the resolution of uncertainty. The growth rate of wealth gets introduced in the stochastic discount factor as a surrogate for tomorrow's utility index. The resulting stochastic discount factor that prices security payoffs is a geometric average of consumption growth and the rate of return on the market portfolio ($R_{M,t+1}$)

$$\langle 5\rangle \qquad m_{t+1} = \beta^{1-\nu}\left(\frac{c_{t+1}}{c_t}\right)^{-\gamma} R_{M,t+1}^{-\nu},$$

where $\gamma \equiv (1-\rho)(1-\nu)$, $\nu \equiv 1-(\alpha/\rho)$; and $\rho < 1$ reflects the investor's attitude towards intertemporal substitution. The relative risk aversion is decreasing in α ($RRA \equiv 1-\alpha$). When $\nu > 1$, $\gamma < 0$ and vice versa. The consumer prefers early resolution of uncertainty when $\nu < 1$ and late resolution otherwise. When $\alpha = \rho$, ν is identically zero, and the consumer is indifferent to the resolution of uncertainty (as in the standard expected utility model).

I.D. Model 4: Ad-hoc factor model

Inspired by the work of Hansen and Jagannathan (1994) and Jagannathan and Wang (1996), we test as our fourth model, the ad-hoc factor model in which the stochastic discount factor incorporates three factors (in addition to the constant):

$$\langle 6\rangle \qquad m_{t+1} = \lambda_0 + \lambda_1\left(\frac{c_{t+1}}{c_t}\right)^{-\gamma} + \lambda_2\left(\frac{c_{t+1}}{c_t}\right)^{-\varrho} R_{M,t+1}^{-\nu} + \lambda_3\left(\frac{c_t}{c_{t-1}}\right)^{-\vartheta}.$$

This specification of the stochastic discount factor includes both the market-based CAPM and consumption growth as special cases. The factor $(c_t/c_{t-1})^{-\vartheta}$ emphasizes the importance of the lagged consumption growth as in a non-separable asset pricing model. The generalization in $\langle 6 \rangle$ can distinguish between competing factor structures by imposing exclusion restrictions on λ_j for $j = 1,2,3$.

The existing research, which has primarily focused on US security market data, has shown that the representative agent paradigm is only partially successful. For instance, Heaton (1995) indicates that habit persistence can explain the equity premium and the risk free rate puzzle but counterfactually suggests a volatile short rate. Cochrane and Hansen (1992) show that the first two moments in the asset pricing data can be explained only when higher volatility in the stochastic discount factor does not manifest itself as a volatile mean of the IMRS. For models without frictions, typically this is a stringent requirement. Hansen *et al.* (1995) and Heaton and Lucas (1992) provide an alternate framework in which the stochastic discount factor incorporates trading frictions and incomplete markets. Nevertheless, we use the frictionless complete markets approach as a starting point to motivate our empirical investigation. This important step should be undertaken before any asset pricing models with incomplete markets and trading frictions relevant to Japanese capital markets are developed and tested.

II. Non-parametric evaluation of asset pricing models

The empirical implementation of the Hansen and Jagannathan (1991) volatility bound test and the Hansen and Jagannathan (1994) specification error test require time series observations on $\{c_{t+1}/c_t,\ R_{M,t+1}\}$ and a vector of test assets. As in Hansen and Singleton (1982) and Ferson and Constantinides (1991), one plus the per-capita growth rate of monthly nondurable and service consumption is used as a proxy for c_{t+1}/c_t. Consistent with common practice, $R_{M,t+1}$ is approximated by the value-weighted real rate of return on Section I of the Tokyo Stock Exchange stocks (e.g. Bakshi *et al.*, 1995; Campbell and Hamao, 1992).

The need for a powerful set of test assets in our non-parameteric tests can never be overstated. A properly chosen vector of test assets, for instance, will result in a robust experimental design on which to base the acceptance or the rejection of competing asset pricing models. One such criteria is that the set of assets be indicative of the persistent time-series and cross-sectional fluctuations in Japanese stock and bond markets. Following Bakshi and Chen (1996), Cecchetti *et al.* (1994), and Hansen and Jagannathan (1994), the test assets must also incorporate conditioning information. This consideration is of practical importance since asset returns contain predictable components. With these goals in mind, the payoff vector includes two primitive assets: (i) the real return on long-term government bond (denoted $RLTGB_t$); and (ii) the value-weighted real returns on Section II of the Tokyo Stock Index stocks (denoted $RTSE2_t$). Synthetic assets are constructed by multiplying each primitive asset return by the lagged value of each primitive asset (i.e. $RLTGB_{t-1}$, $RTSE2_{t-1}$) and by lagged return on the market portfolio ($R_{M,t-1}$). Thus the payoff vector consists of eight assets (six synthetic and two primitive).[3] Finally, for illustrative purposes, the time discount factor β is set equal to 0.99 throughout our empirical investigation. Prefixing the preference parameter will not only reduce the number of parameters that need to be estimated but also facilitate a comparison of the magnitude of the preference parameters from our test methods (i.e. non-parametric versus GMM estimation of each asset pricing model).

The main purpose of the tests in the next two sub-sections is: (i) to provide an understanding of the dimensions along which existing asset-pricing models may be failing in Japan; and (ii) to contrast the results to what we know using US security market data. Given the recent academic interest in the Japanese capital markets, this exercise can offer suggestions on improving the empirical fit if any asset pricing model and will guide the search for better performing stochastic discount factors.

II.A. Hansen–Jagannathan volatility bounds tests

Hansen and Jagannathan (1991) provide a diagnostic test which can be used to judge the usefulness of a stochastic discount factor inherent in any asset pricing model. As demonstrated in their paper, a necessary condition that any asset pricing model must satisfy is that the mean-standard deviation of its stochastic discount factor lie within a region

$$\langle 7\rangle \quad Std(m) \geq Std(x) \equiv \left\{\left[E(\underline{q}) - E(m)E(\underline{x})\right]'\Sigma_x^{-1}\left[E(\underline{q}) - E(m)E(\underline{x})\right]\right\}^{1/2}$$

where $E(m)$ and $Std(m)$ are, respectively, the unconditional mean and standard deviation of the candidate stochastic discount factor; Σ_x is the unconditional variance-covariance matrix of the payoff vector $\underline{x}$, and $\underline{q}$ denotes the vector of asset prices. If the candidate stochastic discount factor does not generate enough variation for any given value of $E(m)$, it will lie outside the Hansen–Jagannathan acceptance region and the asset pricing model can be deemed inconsistent with the asset pricing data.[4]

In Figure 1, the mean-standard deviation pairs for the volatility bound are denoted by the □-curve and correspond to the minimum value of $Std(x)$ for each value of $E(m)$. This curve generates the admissible mean-standard deviation pair that each candidate stochastic discount factor must satisfy. According to the □-curve, the volatility bound is sharp with minimum recorded volatility of 25% which corresponds to the mean stochastic discount factor of 0.9975. Take the stochastic discount factor from time-separable asset pricing model in ⟨3⟩ as an example. The mean-standard deviation pairs for this stochastic discount factor are displayed as the O-curve in Figure 1. Several distinctive patterns emerge. For $\gamma \in (12.0, 12.5)$, the Hansen–Jagannathan (HJ) bounds are not rejected. Initially when γ increases, both the mean and the volatility of the stochastic discount factor decline. But as $|\gamma|$ gets larger, the effect of negative consumption growth rate dominates, which ultimately results

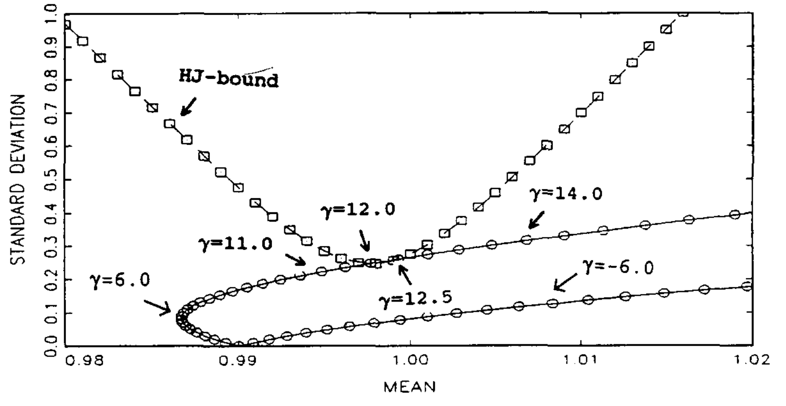

FIGURE 1. Hansen–Jagannathan bounds: time-separable model. *Note*: The Hansen–Jagannathan bounds are illustrated by the □-curve. The O-curve stands for the mean-standard deviation pairs of the candidate IMRS obtained by fixing $\beta = 0.99$ and varying γ.

in a positively sloped O-curve. This switch occurs at $\gamma = 6$. Finally, the minimum magnitude of risk aversion that is required to satisfy the HJ bounds, using Japanese security market data, is 12.0 compared with at least 100 in the United States (e.g. Hansen and Jagannathan, 1991).

By the same token, for the Abel (1990) model (see Figure 2), the γ parameter must lie between 15.50 and 18 for the HJ bounds to be satisfied. For $\gamma < 8.0$, the mean-standard deviation curve is negatively sloped and is positive otherwise. The overall implication is that the inclusion of consumption externality, as in ⟨4⟩, does not result in a better performing stochastic discount factor. Thus when γ is restricted to be less than 15.50, the pricing model in Abel has limited ability to explain the volatility of Japanese asset returns.

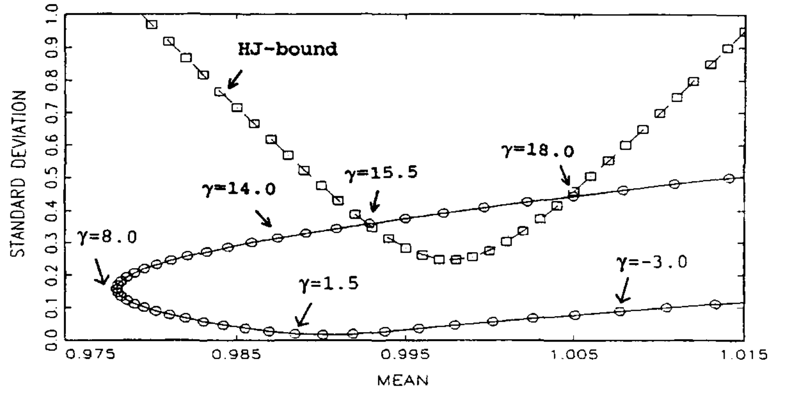

FIGURE 2. Hansen–Jagannathan bounds: Abel's model. *Note*: The Hansen–Jagannathan bounds are illustrated by the □-curve. The O-curve stands for the mean-standard deviation pairs of the candidate IMRS obtained by fixing $\beta = 0.99$ and varying γ.

Ferson and Constantinides (1991) have argued that time non-separability in consumption expenditures increases the variability in the marginal utility of consumption and results in a more volatile IMRS. In Figure 3, the mean-standard deviation pairs for the stochastic discount factor, denoted by the O-curve, are constructed by fixing the value of γ at 13.50 and varying the time non-separable parameter b. The IMRS is constructed following the suggestion in Ferson and Harvey (1992). Note that (i) for $b < 0$ (durability), the mean-standard deviation pairs are outside the volatility bound and both the mean and the volatility of the stochastic discount factor are increasing in $|b|$. But with durability, the incremental increases in the mean and standard deviation of the stochastic discount factor are smaller compared to the habit counterpart (see Figure 3); (ii) increases in γ intensify the effect of habit formation and durability on the volatility of the stochastic discount factor; (iii) since a value of b in the range 0–0.05 (with $\gamma = 13.50$) is required to satisfy the HJ acceptance region, only a modest amount of habit formation is suggested by this exercise; and (iv) When γ is increased to 20 (see the ◇-curve), it has the impact of lowering (increasing) the mean (standard deviation) of the stochastic discount factor. Higher γ induces durability to become consistent with the volatility bounds.

Absent sampling error, the mean-standard deviation frontier for the recursiveutility model is displayed respectively in Figure 4A and B for $\nu = -0.50$ and $\nu = 0.90$ with γ varying. To aid discussion, recall that when $\nu > 1$, then $\gamma < 0$, and vice versa. In Figure 4A, the turning point of the mean-standard deviation frontier for the stochastic discount factor occurs at $\gamma = 4.50$ and $\nu = -0.50$. The HJ bounds are satisfied when ν is fixed at -0.50 and γ varies between 12 and 13. Thus the volatility bounds tests are consistent with values for the elasticity of intertemporal substitution that are in the range 0.115–0.125. Even though the elasticity of intertemporal substitution and risk aversion are

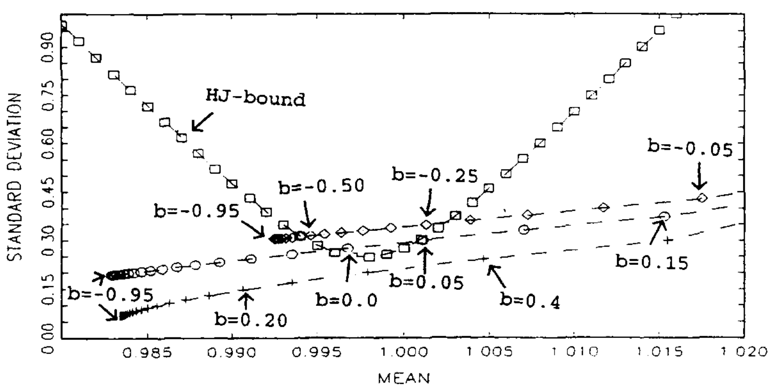

FIGURE 3. Hansen–Jagannathan bounds: time non-separable model. *Note*: The Hansen–Jagannathan bounds are illustrated by the □-curve. The +-curve, the O-curve and the ◇-curve, respectively, stands for the mean-standard deviation pairs of the candidate IMRS obtained by fixing $\gamma = 5$, 13.5, 20 and varying the nonseparable parameter b. In all calculations, $\beta = 0.99$.

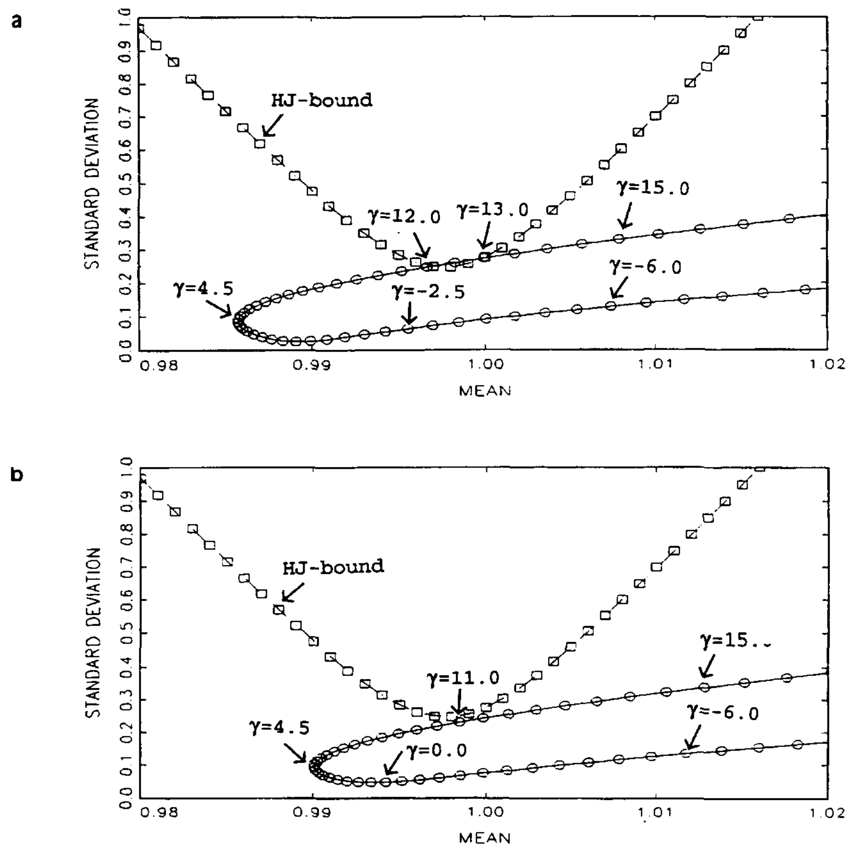

FIGURE 4. Hansen–Jagannathan bounds: Recursive Utility Model. *Note*: The Hansen–Jagannathan bounds are illustrated by the □-curve. The O-curve stands for the mean-standard deviation pairs of the candidate IMRS obtained by fixing $\beta = 0.99$, $\nu = -0.50$ (a) and $\nu = 0.90$ (b), and varying γ.

no longer restricted to be the inverse of each other, the resulting range for the coefficient of relative risk aversion is still large, i.e. 11.50–12.50. These parameter values indicate that early resolution of uncertainty $\nu < 1$ is more consistent with a volatile IMRS.[5] On comparing Figures 4A and B, one can observe that when $|\nu|$ increases, a higher mean-standard deviation pair is attained (holding γ fixed). This means that the ability of the recursive utility model to generate a volatile stochastic discount factor originates from the volatile wealth index rather than consumption growth.

According to Cecchetti *et al.* (1994), two sources of uncertainty emanate when researchers compare the mean–standard deviation pairs from the volatility bound to the stochastic discount factor counterpart. First, the computation of the mean–standard deviation pair for each stochastic discount factor is influenced by the estimated sample moments of the consumption process (e.g. as in expected utility) and the wealth surrogate (as in recursive utility). Second, volatility bounds must be constructed from the asset return data. That is, both the moments of the stochastic discount factor and the volatility bound are data specific and sample dependent. Cecchetti *et al.* (1994) exploit these observations to convert the volatility bound test into a statistical test. Their procedure is particularly appealing when large values of risk aversion result in a noisy (mean) stochastic discount factor (which seriously undermines the ability of the volatility bound diagnostic to discriminate between competing asset pricing models). This desirable property motivates the adoption of their econometric methodology for our empirical analysis.

At the core of the Cecchetti *et al.* (1994) test approach is a statistical assessment of how close $Std(m)$ is to $Std(x)$. The key feature of their econometric technique is that it allows the investigator to reject or accept the hypothesis that a particular class of stochastic discount factor is consistent with the volatility bounds. In their methodology, if the t-value from the one-sided null hypothesis $Std(m) - Std(x) \leq 0$ is less than 1.65 in absolute value, the null is not rejected at the 5% significance level. For further details regarding the asymptotic distribution theory and the exact moment conditions, we refer the reader to their paper. As in Cecchetti *et al.* (1994), the reported standard errors employ the Newey–West (1987a) correction procedure with lag length equal to 11.[6,7]

Panels A through C of Table 1 report the results on the Cecchetti *et al.* (1994) volatility bounds test for the time-separable, the Abel, the time non-separable, and the recursive utility asset pricing models. Displayed for each model are $Std(m)$ and $Std(x)$ for a given $E(m)$ and the t-value for the hypothesis $Std(m) - Std(x) \leq 0$. First consider the benchmark case in Panel A: the time-separable model. Clearly the time-separable model is not rejected when the curvature parameter $\gamma \geq 8.0$. Thecorresponding value absent sampling error is 12.50 (see Figure 1). The main message is that when statistical considerations are taken into account, the mean-standard deviation pairs of the IMRS become consistent with the volatility bounds restriction with reasonable values of the concavity parameter. One can arrive at similar conclusions for the Abel pricing model, where the HJ acceptance region requires that γ lie narrowly between 5 and 6.

Now focus on Panel B of Table 1 which allows the non-separable parameter b to accommodate habit formation and durability. A number of points can be madebased upon this table: (i) Compared to the durability counterpart, both the mean and standard deviation of the IMRS incorporating habit (i.e. $b > 0$) rise faster when γ increases. As an illustration, fix $\gamma = 2.0$. When $b = 0.40$, $Std(m) = 0.137$ and for $b = -0.40$, $Std(m) = 0.028$. (ii) With durability ($b = -0.40$), a curvature magnitude of at least 15.0 is required to maintain consistency with HJ bounds. In contrast, with habit formation, the time non-separable model is not rejected when $\gamma = 2.0$ and $b = 0.40$. Since durability persistently produces less variability in the IMRS whereas habit formation yields more variation, only habit forming preferences can be consistent with the substantial fluctuations observed in Japanese security markets. (iii) The required mean of the IMRS, for which the HJ bounds are satisfied, is close to one.

The results for the recursive utility model reported in Panel C of Table 1 allow γ to vary but ν is fixed at either -0.50 or 0.90. When $\nu > 1$, the mean-standard deviation pairs are quantitatively similar but the restriction $\gamma < 0$ must hold. To save space, we concentrate on the two parameter values. Note that for a given value of γ, *Std*(m) increases as $|\nu|$ increases. On the other hand, for a value of $|\nu|$, changing γ results in only moderate increases in *Std*(m). Therefore, as previously argued, the ability of the recursive utility model to generate a volatile IMRS is driven mostly by the market portfolio

TABLE 1. Volatility bound tests

Panel A: Time-separable and Abel models

	Time-separable				Abel			
γ	$E(m)$	$Std(m)$	$Std(x)$	t-value	$E(m)$	$Std(m)$	$Std(x)$	t-value
0	0.990	0.000	0.474	−5.57	0.989	0.019	0.536	−5.71
1	0.989	0.019	0.537	−5.63	0.989	0.019	0.537	−5.63
2	0.987	0.039	0.588	−5.11	0.989	0.049	0.497	−4.60
3	0.987	0.059	0.624	−4.37	0.992	0.082	0.421	−3.22
4	0.986	0.079	0.642	−3.63	0.995	0.116	0.340	−2.00
5	0.986	0.099	0.643	−2.96	0.999	0.150	0.342	−1.38
6	0.987	0.120	0.626	−2.37	1.004	0.186	0.507	−1.59
7	0.987	0.141	0.591	−1.85	1.010	0.225	0.798	−2.15
8	0.988	0.163	0.540	−1.40	1.018	0.266	1.176	−2.69
9	0.990	0.186	0.475	−1.02	1.027	0.310	1.629	−3.12
10	0.992	0.209	0.403	−0.71	1.037	0.358	2.155	−3.45
15	1.009	0.343	0.724	−0.64	1.109	0.990	5.988	−4.02

Panel B: Time non-separable model

	$b = 0.40$ (Habit)				$b = -0.40$ (Durability)			
γ	$E(m)$	$Std(m)$	$Std(x)$	t-value	$E(m)$	$Std(m)$	$Std(x)$	t-value
0	0.990	0.000	0.477	−5.58	0.990	0.000	0.477	−5.58
1	0.991	0.068	0.454	−4.25	0.988	0.014	0.543	−5.78
2	0.996	0.137	0.328	−1.60	0.987	0.028	0.606	−5.46
3	1.006	0.209	0.591	−1.68	0.986	0.041	0.663	−4.92
4	1.021	0.287	1.338	−2.43	0.985	0.055	0.712	−4.36
5	1.042	0.372	2.429	−2.87	0.984	0.069	0.753	−3.86
6	1.070	0.469	3.885	−3.09	0.983	0.083	0.785	−3.42
7	1.106	0.588	5.783	−3.14	0.983	0.096	0.808	−3.03
8	1.151	0.821	8.188	−1.95	0.982	0.110	0.822	−2.68
9	1.217	1.024	11.69	−2.70	0.982	0.124	0.827	−2.37
10	1.184	4.328	9.921	−0.06	0.982	0.138	0.822	−1.83
15	1.782	3.547	41.61	−1.17	0.986	0.209	0.661	−0.95

— *continued*

TABLE 1 (*Continued*)

Panel C: Recursive utility model

	$\nu = -0.50$				$\nu = 0.90$			
γ	$E(m)$	$Std(m)$	$Std(x)$	t-value	$E(m)$	$Std(m)$	$Std(x)$	t-value
0	0.989	0.025	0.516	−4.14	0.993	0.046	0.360	−1.84
1	0.987	0.032	0.583	−4.18	0.992	0.051	0.403	−1.87
2	0.986	0.046	0.636	−4.04	0.990	0.061	0.442	−1.87
3	0.985	0.063	0.672	−3.74	0.990	0.076	0.471	−1.80
4	0.985	0.082	0.692	−3.33	0.989	0.092	0.486	−1.66
5	0.985	0.101	0.693	−2.88	0.989	0.111	0.485	−1.46
6	0.985	0.121	0.676	−2.41	0.990	0.130	0.469	−1.20
7	0.986	0.142	0.641	−1.95	0.991	0.151	0.439	−1.01
8	0.987	0.163	0.589	−1.52	0.992	0.172	0.398	−0.80
9	0.989	0.186	0.521	−1.13	0.994	0.195	0.354	−0.65
10	0.990	0.209	0.443	−0.80	0.996	0.218	0.324	−0.69
15	1.008	0.341	0.662	−0.54	1.013	0.352	0.925	−0.86

The volatility bound tests reported here are based upon Cecchetti *et al.* (1994). The IMRS being tested are based upon the models outlined in Section II. The asset vector includes eight assets: $RTSE2_t$, $RLTGB_t$ $RTSE2_t \cdot RTSE2_{t-1}$, $RTSE2_t \cdot RLTGB_{t-1}$, $RLTGB_t \cdot RTSE2_{t-1}$, $RLTGB_t \cdot RLTGB_{t-1}$, $RTSE2_t \cdot RTSE1_{1,t-1}$, $RLTGB_t \cdot RTS1_{1,t-1}$, where $RTSE2$, $RLTGB$, and $RTSE1$ denote, respectively, the real returns on the Section II of TSE stocks, long-term government bonds and Section I of TSE stocks. For each estimation, $\beta = 0.99$. The asymptotic standard errors are based on Cecchetti *et al.* (1994, equation ⟨19⟩) and a lag length of 11 is used in the computation of the Newey–West (1987a) covariance matrix. The reported t-value tests the null hypothesis that $Std(m) - Std(x) \leq 0$, where $Std(m)$ and $Std(x)$ are, respectively, the standard deviation of the IMRS and the volatility bound. The critical t-value, above which the null is rejected, is −1.65 at the 5% and −2.33 at the 1% significance level.

rather than consumption growth. Furthermore adding non-zero values of $|\nu|$ is consistent with the volatility bounds. For instance with $\nu = 0.90$, the HJ volatility bounds are not rejected when $\gamma > 5$. Similarly, by fixing $\nu = -0.50$, the recursive utility asset pricing model is not rejected when $\gamma > 8.0$. Recall from Figure 3 that when $\nu = -0.50$, the corresponding γ value, absent statistical considerations, is 12.0. Finally, when $\gamma \equiv 0$, $m_{t+1} = (\beta^{1-\nu}/R^{\nu}_{M,t+1})$, the returns on the market portfolio are sufficient to discount future payoffs. For this model, the volatility bounds are always violated since the absolute t-values are in excess of 1.65.

Despite the formal econometric characterization of the rejection procedure, small sample properties of the Cecchetti *et al.* (1994) estimators are virtually unknown. For example it is possible that the *power* of the t-test may be lower in some models than in others. Therefore, caution must be exercised when evaluating the t-ratios across the pricing models and in attempting to infer which one fits the security market data better. This observation motivates the adoption of the Hansen and Jagannathan (1994) specification-error tests which are considered next.

II.B. Hansen–Jagannathan specification error tests

Clearly, not all stochastic discount factors will price every asset payoff correctly. The specification-error of Hansen and Jagannathan (1994) calculates the maximum pricing error associated with a stochastic discount factor and measures the least square distance between a candidate stochastic discount factor (denoted by y) and the set of admissible stochastic discount factors (denoted by M). The specification error, δ, is obtained as a solution to the following minimization problem:

$$\langle 8 \rangle \qquad \delta^2 = \min_{m^* \in M} E\left[(y - m^*)^2\right].$$

Hansen *et al.* (1995) demonstrate that for the special case $y \equiv 0$, the minimization problem in $\langle 8 \rangle$ is equivalent to the volatility bound in $\langle 7 \rangle$. According to Hansen and Jagannathan (1994), the specification-error criterion is[8]

$$\langle 9 \rangle \qquad \delta = \left\{\left[E(\underline{q}) - E(m\underline{x})\right]' E(\underline{x}\underline{x}')^{-1}\left[E(\underline{q}) - E(m\underline{x})\right]\right\}^{1/2}$$

All the variables are as defined earlier. The distance criterion for any admissible stochastic discount factor that correctly prices the set of payoffs under investigation is identically zero. Among the set of stochastic discount factors, the one with the smallest distance measure is considered to be the best.

In addressing the statistical failure of intertemporal asset pricing models, the specification-error test of Hansen–Jagannathan (1994) has quite a few appealing properties. (a) The minimized specification-error criteria permit the researcher to distinguish between alternate asset pricing specifications and in particular allow one to draw conclusions on the empirical performance of non-nested pricing models. Thus, for example, if the specification-error from the time separable asset pricing model is *lower* compared to the recursive utility model, the data are rejecting the recursive utility asset pricing model in favor of the alternative. This feature of model assessment substitutes the corresponding shortcoming of the volatility bounds test. (b) The finite-sample properties of the statistic in $\langle 9 \rangle$ are better when a fixed $E\,(x\,x')^{-1}$ matrix is used instead of the weighting matrix in Hansen (1982) and Newey and West (1987a). (See Cochrane, 1994 for a further discussion of this point.) (c) When $y \equiv 0$, the minimum distance problem in $\langle 9 \rangle$ reduces to finding the lower volatility bound. This means that compared to the volatility bounds test, the specification-error test imposes a stronger set of conditions that any asset pricing model must satisfy. Specifically the preference parameters satisfying the volatility test are less robust compared to the specification-error analogues. (d) Finally, the minimum distance measure tests can be implemented even in the presence of market frictions (e.g. short-sales constraints and transaction costs).

The distance measure, δ, for the different stochastic discount factors is displayed in Panel A of Table 2. The standard errors, reported in parentheses, are calculated using Proposition 2.2 in Hansen *et al.* (1995) and uses the Newey–West (1987) correction with lag length equal to 11. For the given range of γ, the maximal pricing errors from the Abel model are smaller in magnitude than the time-separable model except when $\gamma = 0.0$ or 15. Thus, relative to the time-separable asset pricing model, the sample evidence against consumption externality is smaller. Compared to the time-separable model, the pricing errors are consistently smaller when the IMRS incorporates time non-separability. For instance when $\gamma = 2.0$, the distance measure δ for the time-separable model, habit formation ($b = 0.40$), and durability ($b = -0.40$), respectively, equals 0.245, 0.235, and 0.241. Even though the recursive utility model is observationally equivalent to the time-separable model when $\nu \equiv 0$, introducing either positive or negative values of ν does not decrease the specification-error but instead increases it. For $\gamma < 10$, the distance measure for the recursive utility model is persistently larger compared to the counterpart from the time non-separable model.

We can also apply the specification-error test when the stochastic discount factor associated with an asset pricing model depends upon a vector of unknown preference parameters. For any non-nested class of asset pricing models, Hansen and Jagannathan (1994) and Hansen *et al.* (1995) establish that the specification-error can be used as a goodness-of-fit measure for choosing between competing models. They suggest a two-step procedure. *Step 1*: Minimize the specification measure (with the prefixed weighting matrix) in equation $\langle 9 \rangle$ by choosing the vector of unknown parameters. *Step 2*: Calculate the specification-error corresponding to these parameter values. The stochastic discount factor with the lowest distance measure is the least misspecified in the sample. Panel B of Table 2 reports the results from this exercise. Among our parameterized models, the stochastic discount factors for the habit forming

TABLE 2. Hansen–Jagannathan specification error tests

γ	$b = 0$	Abel	$b = 0.40$	$b = -0.40$	$\nu = -0.50$	$\nu = 0.90$
			Panel A: Specification errors			
0	0.247	0.248	0.242	0.242	0.247	0.248
	(0.069)	(0.069)	(0.068)	(0.068)	(0.069)	(0.070)
2	0.245	0.243	0.235	0.241	0.245	0.247
	(0.070)	(0.070)	(0.068)	(0.067)	(0.070)	(0.071)
5	0.243	0.237	0.230	0.240	0.244	0.245
	(0.072)	(0.073)	(0.069)	(0.067)	(0.072)	(0.073)
10	0.243	0.236	0.284	0.238	0.243	0.245
	(0.076)	(0.079)	(0.077)	(0.066)	(0.076)	(0.077)
15	0.247	0.265	0.578	0.238	0.246	0.250
	(0.082)	(0.092)	(0.219)	(0.066)	(0.081)	(0.083)

— *continued*

TABLE 2 (*Continued*)

Panel B: Minimized specification error

Model	Minimized δ	Parameters
Time	0.24727	$\gamma = 8.40$
Separable	(0.070)	
Abel	0.23452	$\gamma = 8.10$
	(0.076)	
Time	0.22841	$\gamma = 2.50$
Nonseparable	(0.069)	$b = 0.55$
Recursive	0.24265	$\gamma = 8.5$
Utility	(0.074)	$\nu = -0.17$
Ad Hoc	0.18050	$\lambda_0 = 31.06$
Factor	(0.329)	$\lambda_1 = 26.65$
		$\lambda_2 = -2.95$
		$\lambda_3 = -53.83$
$\lambda_2 \equiv 0$	0.23523	$\lambda_0 = -20.64$
$\lambda_3 \equiv 0$	(0.087)	$\lambda_1 = 21.67$
$\lambda_1 \equiv 0$	0.24575	$\lambda_0 = 0.85$
$\lambda_3 \equiv 0$	(0.068)	$\lambda_2 = 0.14$
$\lambda_1 \equiv 0$	0.19937	$\lambda_0 = 51.51$
$\lambda_2 \equiv 0$	(0.210)	$\lambda_1 = -50.60$
$\lambda_3 \equiv 0$	0.23100	$\lambda_0 = -27.91$
	(0.098)	$\lambda_1 = 30.58$
		$\lambda_2 = -1.64$

Estimation of the specification error, δ, is based on Hansen and Jagannathan (1994, equation 2.10). The standard errors, reported in parentheses, are estimated following Hansen *et al.* (1995, Proposition 2.2). A lag length of 11 is employed for the Newey–West (1987a) correction. The payoff vector includes eight assets: $RTSE2_t$, $RLTGB_t$, $RTSE2_t \cdot RTSE2_{t-1}$, $RTSE2_t \cdot RLTGB_{t-1}$, $RLTGB_t \cdot RTSE2_{t-1}$, $RLTGB_t \cdot RLTGB_{t-1}$, $RTSE2_t \cdot RTSE1_{1,t-1}$, $RLTGB_t \cdot RTS1_{1,t-1}$, where $RTSE2$, $RLTGB$, and $RTSE1$ denote, respectively, the real returns on Section II of TSE stocks, long-term government bonds and Section I of TSE stocks. For each estimation, set $\beta = 0.99$. The reported minimum δ in Panel B is obtained by choosing the preference parameters of various models to minimize the Hansen–Jagannathan specification error (see eq. ⟨9⟩).

preference in ⟨2⟩ provides the best characterization of Japanese security prices. The time non-separable model has a minimum specification error of 0.22841 and this minimum occurs at $(\gamma, b) = (2.50, 0.55)$. This is consistent with high values of habit persistence in the Japanese consumption data. While presenting international evidence on the time non-separable pricing model, Braun *et al.* (1993, Tables 2 and 3) record $b \in (0.64, 0.75)$, a magnitude similar to the minimized distance criteria here. Similarly, as reported, the minimum specification-error for the time-separable and Abel models are, respectively, 0.24727 and 0.23452. This minimum distance occurs at γ values of 8.40 and 8.10, respectively. Somewhat surprisingly, the minimized distance measure for the recursive utility model departs only slightly from the time-separable model. The minimized distance is obtained at (γ, ν) values of (8.50, -0.17). Based upon the distance criteria, therefore, the IMRS of the recursive utility model does not provide a significant performance improvement.

The ad-hoc factor model yields further insights into the relative performance of each asset pricing model. While estimating the general ad-hoc model in ⟨6⟩, we encountered significant numerical difficulties in obtaining the minimized distance criteria for the choice of γ, ϱ, ν, and ϑ. To focus on the performance issues, from now on, specialize the ad-hoc model in ⟨6⟩ to:

$$\langle 10\rangle \qquad m_{t+1} = \lambda_0 + \lambda_1\left(\frac{c_{t+1}}{c_t}\right)^{-1} + \lambda_2\left(\frac{c_{t+1}}{c_t}\right)^{-1} R_{M,t+1}^{-1} + \lambda_3\left(\frac{c_t}{c_{t-1}}\right)^{-1}.$$

This asset pricing model obtains when $\gamma = \varrho = \nu = \vartheta \equiv 1$. The IMRS in ⟨10⟩ can be justified theoretically in economies with logarithmic utility function or by taking an appropriate Taylor series expansion of the stochastic discount factor. Results obtained from conducting this exercise are reported in the last row of Panel B of Table 2. When the minimized distance is 0.1805, the estimated parameter vector is: (31.06, 26.65, -2.95, -53.83). For the class of stochastic discount factors under study, the minimized distance from the ad-hoc (linear) model is modest (compared to competing models). But the standard error on the distance measure is substantial, which renders ambiguous any deductions based on specification-error. One can also test the effectiveness of each ad-hoc factor by imposing exclusion restrictions in ⟨10⟩ and by contrasting the associated minimized specification-error. When $\lambda_2 = \lambda_3 = 0$, for example, the ad-hoc stochastic discount factor is linearly related to the inverse of consumption growth. In this case, the minimized specification-error is 0.23523. Now restrict $\lambda_1 = \lambda_3 = 0$. The (constraint) minimized distance occurs at 0.2457, which is higher compared to the counterpart from consumption growth. In particular, when the stochastic discount factor comprises of lagged inverse consumption growth, the fit of the ad-hoc factor model improves substantially and reduces the minimized specification error. Lagged consumption growth, however, tends to increase the standard error on the minimum specification-error, which generates unreliable statistical inferences.

It is somewhat disconcerting that the minimized distance measures for the parameterized models are in the same narrow range. One can attribute this observation to either: (i) the inability of the current set of test assets to discriminate between the asset pricing models, resulting in a suboptimal experimental design; or (ii) the Hansen–Jagannathan distance measure being ineffective. The latter issue is beyond the scope of the paper. To address concerns regarding experimental design, we included the time-t long-term Japanese corporate bond returns to our set of test assets and found that the ordered ranking of the minimized distance criteria remained unchanged. This means that our conclusions regarding the relative performance of the asset pricing models is robust.

Before closing this section, it remains to be emphasized that recursive preferences have performed poorly relative to other preference structures. This is despite the fact that under this preference ordering: (a) the risk premium is determined by both its covariance with consumption growth and the rate of return on the market portfolio; and (b) the risk aversion and the elasticity of intertemporal substitution are no longer reciprocals. The estimated parameters from the volatility bounds test and the specification error test provide evidence that early resolution of uncertainty is empirically relevant. The GMM-based tests in Epstein and Zin (1991), however, found evidence to the contrary (using US data). The observation that Japanese consumer–investors prefer early resolution of uncertainty can be useful in the context of investment policy and portfolio management. As suggested by Epstein and Zin (1989), early resolution can improve planning. For further discussion of this point, see the treatment in Epstein and Zin (1989, 1991).

III. GMM tests of the Euler equation

This section applies Hansen's (1982) generalized method of moments to test the Euler equation restrictions implied by each asset pricing model. The objective is twofold. One goal is to determine whether the magnitude of the estimated preference parameters are consistent with the non-parametric tests. The other goal is to use the GMM criterion function to conduct goodness-of-fit test and likelihood ratio-based exclusion tests of nested asset pricing models. When the restrictions implied by the pricing equation in $\langle 1 \rangle$ are true, then under the null hypothesis $E(\varepsilon_{t+1} \otimes Z_t) = 0$. Here $\varepsilon_{t+1} \equiv m_{t+1} R_{t+1} - 1$ and Z_t is the set of conditioning information. Let G_T denote the sample counterpart of $\varepsilon_t \otimes Z_t$. The GMM estimation, for sample size T, is based upon minimizing the quadratic form $G_T{}' W_T G_T$, where W_T is an optimal weighting matrix in the sense of Hansen (1982). As demonstrated in Hansen (1982), T times the minimized GMM criterion function, the J_T-statistic, is χ^2-distributed with degrees of freedom equal to the number of orthogonality conditions minus the number of parameters to be estimated. A high value of the GMM criterion function indicates a misspecified asset pricing model. In Ferson and Foerster (1994), they show in a small sample that a more accurate GMM test statistic can be constructed when the GMM criterion function is iterated to achieve a prespecified convergence criteria. To enhance the small sample properties of our estimators, the implementation of the GMM uses the iterated GMM approach (see also Hansen *et al.*, 1994).

The choice of information variables to be included in the GMM-based tests is guided by previous empirical work and constructed using financial variables that capture return dynamics (see Ferson and Constantinides, 1991; Bakshi and Chen, 1996). The instruments include a constant and two lags each of the nominal returns on Section I of Tokyo Stock Exchange stocks, the default premium, the term premium, and the nominal genasaki (interest) rate.[9] The set of assets to be jointly estimated via GMM is restricted to two to keep the number of orthogonality conditions to a reasonable number (< 20). To provide powerful tests of the overidentifying restrictions, the GMM tests include the assets $RET1$ and $RET2$ which are constructed using conditioning information (for details, see the Appendix).

Panel A of Table 3 reports the estimation results of the parameters $\{\gamma, b\}$ for the time non-separable asset pricing model and Abel's (1990) model. The

TABLE 3. GMM tests of the Euler equation

Panel A: Non-separable and Abel models						
	Time non-separable				Abel	
Assets	γ	b	J_T	A_T	γ	J_T
RSMALL	1.82	−0.53	28.68	2.32		
and RET1	(1.35)	(1.95)	[0.03]	[0.13]		
	[0.18]	[0.78]	{16}			
	1.09	$b \equiv 0$	28.98		0.96	29.31
	(0.99)		[0.03]		(0.75)	[0.03]
	[0.27]		{17}		[0.20]	{17}
RSMALL	1.89	−0.33	28.55	2.03		
and RET2	(1.38)	(0.86)	[0.03]	[0.15]		
	[0.17]	[0.70]	{16}			
	1.24	$b \equiv 0$	28.82		1.06	29.18
	(1.01)		[0.04]		(0.74)	[0.03]
	[0.22]		{17}		[0.15]	{17}
RSMALL	1.82	−0.41	29.07	2.35		
and RTSE2	(1.36)	(1.20)	[0.02]	[0.13]		
	[0.18]	[0.74]	{16}			
	1.15	$b \equiv 0$	29.34		1.02	26.64
	(1.00)		[0.03]		(0.75)	[0.03]
	[0.25]		{17}		[0.18]	{17}

Estimation of the following Euler equation is based on Hansen's (1982) generalized method of moments,

$$E_t(m_{t+1}R_{j,t+1}|Z_t) = 1,$$

where Z_t contains a constant and two lags each of term premium, default premium, the nominal returns on the Section I of TSE stocks, and the nominal interest rate. The standard errors reported in parentheses are based on the simple covariance matrix estimator outlined in Hansen (1982). The p-value in brackets tests the hypothesis that the estimated parameter equals zero. The degree of freedom df (reported in curly brackets) is the number of moment conditions minus the number of parameters to be estimated. The J_T-statistic tests whether the overidentifying restricions of the model are true with the degrees of freedom, df. The statistic, A_T, which is asymptotically $\chi^2(1)$-distributed, tests the exclusion restriction, $b = 0$. For each estimation, set $\beta = 0.99$. $RTSE2$ is the real return on Section II of TSE stocks, RSMALL is the real return on a portfolio of small stocks and RET1, RET2 are synthetic assets constructed from conditioning information. Details on the construction of these variables is in the Appendix. The reported estimation for $b \equiv 0$ implies the time separable model.

— *continued*

TABLE 3 (*Continued*)

Panel B: GMM tests of the seasonal non-separable model

Assets	γ	b	J_T	A_T
RSMALL	−0.15	−0.11	73.67	2.32
RET1	(0.04)	(0.20)	[0.01]	[0.13]
	[0.00]	[0.59]	{48}	
RSMALL	−0.15	−0.10	70.12	2.03
RET2	(0.04)	(0.20)	[0.02]	[0.15]
	[0.00]	[0.62]	{48}	
RSMALL	−0.15	−0.11	72.98	2.35
RTSE2	(0.04)	(0.20)	[0.01]	[0.13]
	[0.00]	[0.59]	{48}	

Estimation of the following Euler equation is based on Hansen's (1982) generalized method of moments,

$$\varepsilon_{t+1} = \beta R_{t+1}\left\{\left[\frac{c_{t+1} - bc_{t-11}}{c_t - bc_{t-12}}\right] - b\beta^{12}\left[\frac{c_{t+13} - bc_{t+1}}{c_t - bc_{t-12}}\right]^{-\gamma}\right\}$$

$$+ b\beta^{12}\left[\frac{c_{t+12} - bc_t}{c_t - bc_{t-12}}\right]^{-\gamma} - 1. \ \varepsilon_{t+1} \sim MA(12).$$

The information variables in the set Z_t contain a constant and 12 lags each of term premium, default premium, the nominal returns on the Section I of TSE stocks,the nominal interest rate. The standard errors reported in parentheses are based on the simple covariance matrix estimator as outlined in Hansen (1982). The p-value in brackets tests the hypothesis that the estimated parameter equals zero. The degree of freedom df (reported in curly brackets) is the number of moment conditions minus the number of parameters to be estimated. The J_T-statistic tests whether the overidentifying restrictions of the model are true with the degrees of freedom, df. The statistic, A_T, which is asymptotically $\chi^2(1)$-distributed, tests the exclusion restriction, $b = 0$. For each estimation, set $\beta = 0.99$.

benchmark case $b \equiv 0$ corresponds to the standard expected utility model. In that case, the estimate of the risk aversion parameter γ varies between 1.09 and 1.24 and is statistically insignificant. The standard errors, reported in parentheses, are calculated using the simple covariance matrix outlined in Hansen (1982). The goodness-of-fit J_T-statistics range between 28.34 and 28.98, the highest p-value being 3%. The time-separable model is thus rejected using GMM estimation. When the consumption externality enters the IMRS, the estimates of γ are now close to 1.0 but statistically not different from zero. Again, the goodness-of-fit criteria are large and the pricing model in Abel (1990) can be rejected.

The empirical investigation of Ferson and Constantinides (1991) finds that in the US, habit persistence dominates the effect of durability.[10] However, GMM tests conducted using the monthly Japanese stock and bond market data are not so supportive: the estimate of the time non-separable parameter b are negative and range between -0.33 and -0.53 with a minimum p-value of 0.70. The concavity parameter γ ranges from 1.82 to 1.89 and is also statistically insignificant. We are thus led to conclude that *neither* durability nor habit formation is supportable in the seasonally adjusted monthly data and the standard errors on b are too high to place any reliance on the point estimates. The p-values for the 'goodness of fit' J-statistics are less than 5%, and based upon the J-statistic the time non-separable model can be rejected. Recall that the time-separable asset pricing model nests within the time non-separable model. To provide some documentation that introducing non-zero values of b is preferred from the standpoint of statistical significance, we conduct likelihood ratio-based hypothesis tests by constructing:

$$A_T = TG'_{T,U}W_{T,U}G_{T,U} - TG'_{T,R}W_{T,U}G_{T,R}. \qquad \langle 11 \rangle$$

This statistic can assess the relative fit of the two asset pricing models (separable vs non-separable). In the first step, estimate the unrestricted asset pricing model and compute the GMM criterion function. Next impose the restriction $b = 0$ and recompute the criterion function (using the weighting matrix from the unrestricted model). The resulting A_T-statistic is $\chi^2(1)$-distributed (e.g. Newey and West, 1987b). The reported p-values (Panel A of Table 3) indicate that introducing non-separability in consumption to the asset pricing framework does not improve its ability to fit Japanese security market data.

By way of contrast, Braun *et al.* (1993) obtain: (i) a positive non-separability parameter (see their Tables 2 and 3) using quarterly Japanese consumption and asset pricing data; and (ii) no rejections of the J-statistics, which tends to support the habit formation model. These potentially divergent findings can be reconciled by noting differences in (i) observation frequency (monthly vs quarterly); and (ii) the set of conditioning variables and test assets. After subjecting the non-separable model to alternate sets of test assets and information variables, we found that to be a non-issue (see Figure 5 and the discussion to follow). In particular, the estimated sign (and range) of b is similar, and in all cases the non-separable parameter is statistically insignificant. We nonetheless focus attention now on the technical difficulties which may hinder empirical support for the non-separable asset pricing model through the three controlled experiments below:

1. According to Ferson and Constantinides (1991) and Braun *et al.* (1993), the GMM criterion function is nonlinearly related to the non-separable parameter. This nonlinear association can induce large asymptotic standard errors and result in a *local* rather than a *global* minimum for the GMM criterion function. Further when both durability and habit persistence are present in consumption data, the non-separable parameter (with a single lag) does not have enough degrees of freedom to capture both effects. To address these complications, we prefix $b \in (-0.9, 0.9)$ to minimize the GMM criterion function by the choice of γ. We continue to maintain $\varepsilon_t \sim MA(1)$. The sensitivity of the GMM criterion function to the non-separability parameter, for different sets of information variables, is displayed in the +-curve, the □-curve and the O-curve in Figure 5. This

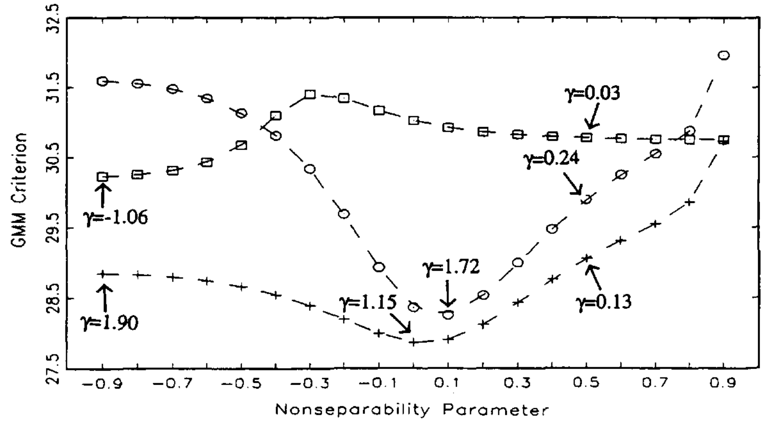

FIGURE 5. GMM criterion functions. *Note*: The +-curve, the □-curve, and the O-curve respectively plot the GMM criterion function for the set of conditioning variables INST1, INST2, and INST3 when the non-separable parameter b varies. The estimations are based upon monthly data. The set of conditioning variables INST1 consists of a constant and two lags each of NTSE1, NINT, TERM, and DEF. The set of conditioning variables INST3 consists of a constant and two lags each of TERM and dividend-yield. The □-curve has been scaled by 1.70.

experiment also illustrates the robustness of our conclusions to the choice of conditioning information. Let us focus on the +-curve and the □-curve, which are constructed using instruments that include return dynamics and the dividend-yield. When $b < 0$, the GMM criterion function curve is negatively sloped in relation to the non-separable parameter with $b = 0.00$ ($b = 0.10$) and $\gamma = 1.15$ ($\gamma = 1.72$) being the (approximate) turning point of the +(□)-curve. The minimized GMM criterion function value is 27.86 (28.73). These results are consistent with the previous conclusions from GMM estimation and with the Hansen–Jagannathan (1991) volatility bound tests. When consumption growth is included in the conditioning information (see the O-curve), the data supports durability. The overall implication is that with financial instruments only a modest amount of habit persistence is consistent with monthly Japanese consumption expenditures data.

2. Ferson and Constantinides (1991) show that the time-series properties of consumption and half-lives of durability and habit are related to the observation frequency. To mitigate the effect of time aggregation on the estimated nonseparable parameter, we aggregate monthly consumption expenditures to quarterly frequency (see Porter and Wheatley, 1995). This yields 80 quarterly observations (for each variable) and allows us to examine the Euler equation restriction implied in ⟨1⟩ using quarterly data. The results (not reported) indicate that: (i) the reported values of b are positive with estimated magnitude indistinguishable from zero. The GMM estimation with quarterly data thus admits *neither* durability nor habit

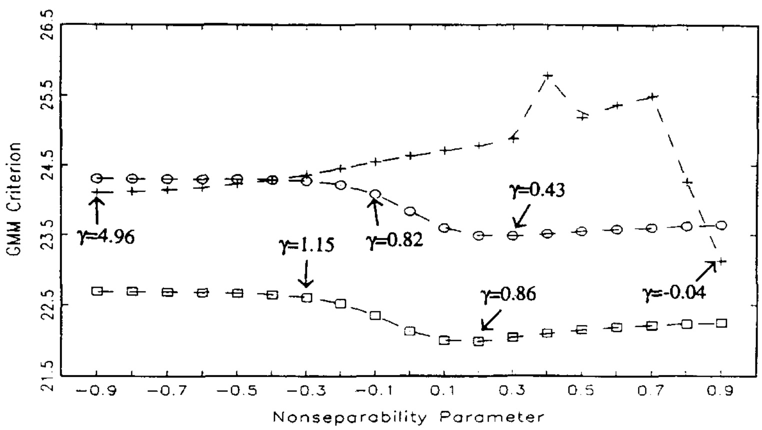

FIGURE 6. GMM criterion functions. *Note*: The +-curve, the □-curve, and the O-curve, respectively, plot the GMM criterion function for the set of conditioning variables INST1, INST2, and INST3 when the non-separable parameter b varies. The estimations are based upon quarterly data. The set of conditioning variables INST1 consists of a constant and two lags each of NTSE1, NINT, TERM, and DERF. The set of conditioning variables INST2 consists of a constant and one lag each of TERM and dividend-yield. The □-curve and the O-curve have been scaled by 1.70 and 2.0, respectively.

formation; (ii) the estimated γ is often less than one; and (iii) the time-nonseparable asset pricing model can be statistically rejected. Since the results from this exercise are similar to that from monthly data, repetition is avoided here. In Figure 6, the +-curve, the □-curve and the O-curve again plot the sensitivity of the criterion function to the non-separable parameter. Notice from Figure 6 that: (i) the minimized value of the criterion function is *not* invariant to the conditioning information; (ii) the estimated risk aversion parameters consistent with minimized criterion function are now smaller compared to the monthly counterpart (range from 0 to 0.86); and (iii) the effect of habit formation dominates the effect of durability for quarterly consumption data. For the three curves the non-separability parameter is respectively: 0.90, 0.20, and 0.15. This is a strong departure from the monthly data where *neither* durability nor habit formation appeared to be dominant. The empirical results lend support to the claim in Ferson and Constantinides (1991) that habit forming preferences are empirically relevant at lower frequencies.

3. One further possibility for the relatively poor fit of the non-separable model and the resulting imprecise estimates is that seasonal adjustments in consumption induce spurious correlations. Of special relevance here is the work of Ferson and Harvey (1992) who show that a seasonal non-separable model in which the subsistence level of consumption depends on the previous season is empirically relevant. Motivated by their paper, we assume a timing convention where consumption decisions are monthly and subsistence level of consumption depends upon the consumption expendi-

tures in the same month of last year. The results are displayed in Panel B of Table 3. Again the point estimates of the seasonal non-separable parameter b are negative and statistically insignificant. The concavity parameter is numerically close to zero and in all cases not statistically significant. The goodness-of-fit tests are rejected with the highest p-value of 2%. Thus we can arrive at the conclusion that seasonal adjustments in consumption are not the underlying cause of the rejection of the time non-separable model.

Now concentrate on Panel A of Table 4 which reports the estimation of the recursive utility model. The estimates of γ range from 0.55 to 0.62 and those of ν, between 0.80 and 0.82. The implied values of the parameter pair (ρ, α) are hence close to $(-1.75, -0.35)$. The estimated parameters indicate that the relative risk aversion $RRA = 1.35$ and the elasticity of intertemporal substitution $(1/(1\text{-}\rho))$ is approximately 0.36. These estimates are analogous to Epstein and Zin (1991), for US data, who obtained: (i) coefficient of relative risk aversion that did not differ substantially from one; and (ii) elasticity of intertemporal substitution less than one. Note that since a constant term is included as an instrument variable in the GMM estimation, the comparison of the conditional vs unconditional restrictions are difficult to substantiate. Nevertheless, the estimated elasticity of intertemporal substitution (risk aversion) is similar (lower) in magnitude compared to the counterpart from the non-parameteric tests. Embedded in ⟨5⟩ is the restriction $\nu \equiv 0$ (i.e. $(\alpha/\rho) \equiv 1$), in which case the recursive utility model reduces to the standard expected utility paradigm. Following the sequence of steps used to compute ⟨11⟩, we can test the restrictions nested within the recursive utility model. The A_T-statistic in the last column, for example, tests the exclusion restriction: $\nu \equiv 0$. The low p-values for this test indicate that relative to the standard expected utility model, the recursive preferences have incremental explanatory power in explaining Japanese stock market behavior. Based upon the goodness-of-fit

TABLE 4. GMM Tests of the Euler Equation

Panel A: Recursive utility				
Assets	γ	ν	J_T	A_T
RSMALL RET1	0.55 (0.78) [0.47]	0.80 (0.17) [0.00]	25.28 [0.07] {16}	18.99 [0.00]
RSMALL RET2	0.62 (0.78) [0.43]	0.82 (0.17) [0.00]	25.12 [0.07] {16}	20.00 [0.07]
RSMALL RTSE2	0.59 (0.78) [0.45]	0.81 (0.17) [0.00]	25.30 [0.06] {16}	19.30 [0.00]

— *continued*

TABLE 4 (*Continued*)

Panel B: Ad-hoc factor model

Assets	λ_0	λ_1	λ_2	λ_3	J_T	A_T
RSMALL	0.04	0.27	0.76	−0.06	24.51	
and RET1	(1.19)	(0.99)	(0.20)	(0.68)	[0.04]	
	[0.98]	[0.79]	[0.00]	[0.93)	{14}	
	−0.03	0.27	0.76		24.51	0.01
	(0.91)	(0.99)	(0.20)		[0.06]	[0.93]
	[0.97]	[0.78]	[0.00]		{15}	
	−0.53	2.02		−0.50	36.27	14.06
	(1.18)	(0.88)		(0.68)	[0.00]	[0.00]
	[0.65]	[0.02]		[0.46]	{15}	
RSMALL	−0.06	0.29	0.78	−0.01	24.32	
and RET2	(1.16)	(0.98)	(0.20)	(0.68)	[0.04]	
	[0.96]	[0.77]	[0.00]	[0.99)	{14}	
	−0.07	0.29	0.78		24.32	0.00
	(0.91)	(0.98)	(0.20)		[0.06]	[0.99]
	[0.94]	[0.77]	[0.00]		{15}	
	−0.69	2.05		−0.37	36.12	14.80
	(1.15)	(0.87)		(0.68)	[0.00]	[0.00]
	[0.55]	[0.02]		[0.59]	{15}	
RSMALL	−0.05	0.32	0.77	−0.03	24.39	
RTSE2	(1.19)	(0.99)	(0.20)	(0.67)	[0.04]	
	[0.96]	[0.75]	[0.00]	[0.96)	{14}	
	−0.09	0.32	0.77		24.39	0.00
	(0.91)	(0.99)	(0.20)		[0.06]	[0.96]
	[0.92]	[0.75]	[0.00]		{15}	
	−0.64	2.03		−0.40	36.51	14.12
	(1.18)	(0.88)		(0.68)	[0.00]	[0.00]
	[0.59]	[0.02]		[0.56]	{15}	

Estimation of the following Euler equation is based on Hansen's (1982) generalized method of moments,

$$E_t(m_{t+1}R_{j,t+1}|Z_t) = 1,$$

where Z_t contains a constant and two lags each of term premium, default premium, the nominal returns on the Section I of TSE stocks, and the nominal interest rate. The standard errors reported in parentheses are based on the simple covariance matrix estimator outlined in Hansen (1982). The p-value in brackets tests the hypothesis that the estimated parameter equals zero. The degree of freedom df (reported in curly brackets) is the number of moment conditions minus the number of parameters to be estimated. The J_T-statistic tests whether the over-identifying restrictions of the model are true with the degrees of freedom, df. The statistic, A_T, which is asymptotically $\chi^2(1)$-distributed tests the exclusion restriction, $\nu = 0$ (for the recursive utility model). For each estimation, set $\beta = 0.99$. $RTSE2$ is the real return on Section II of TSE stocks, RSMALL is the real return on a portfolio of small stocks and RET1, RET2 are synthetic assets constructed from conditioning information. Details on the construction of these variables is in the Appendix.

J-statistics, the recursive utility asset pricing model is not rejected. The highest *p*-value is 7% and the lowest GMM criterion function value is 25.12.

The results from the GMM estimation of the ad-hoc factor model in ⟨10⟩ can be interpreted as follows (see Panel B of Table 4): (i) λ_2 is statistically significant (at the 5% level) with a magnitude in the vicinity of 0.76–0.78; (ii) the GMM criterion function are large and based upon the *p*-value, the overidentifying restrictions imposed by the parameterized ad-hoc factor model are rejected; and (iii) both (inverse) lagged consumption growth and contemporaneous (inverse) consumption growth receive less weight in the data as demonstrated via the A_T statistics, which jointly test the exclusion restrictions $\lambda_1 = 0$ and $\lambda_3 = 0$. Overall, the ad hoc model where consumption growth and the market return enter multiplicatively provides a decent characterization of the Japanese asset pricing data.

Our final econometric methodology exploits the informal diagnostic in Ferson and Constantinides (1991). To gauge the implication of each pricing model (non-separable, Abel, and recursive utility) for the Japanese equity premium, add a parameter θ to the pricing condition

$$\langle 12 \rangle \qquad E_t\{m_{t+1}(R_{i,t+1} + \theta)\} = 1,$$

TABLE 5. Estimation of pricing errors

	Panel A: Various models						
	Time non-separable				Abel		
Assets	γ	b	θ	J_T	γ	$\tilde{\theta}$	J_T
RSMALL	3.48	−0.89	0.006	26.74			
RET1	(1.83)	(6.23)	(0.004)	[0.06]			
	[0.06]	[0.89]	[0.16]	{15}			
	1.80	$b \equiv 0$	0.005	25.67	0.98	0.003	28.83
	(1.19)		(0.004)	[0.06]	(0.76)	(0.003)	[0.03]
	[0.13]		[0.25]	{16}	[0.19]	[0.40]	{16}
RSMALL	3.71	−0.55	0.007	26.53			
RET2	(1.84)	(1.04)	(0.005)	[0.03]			
	[0.07]	[0.60]	[0.15]	{16}			
	1.89	$b \equiv 0$	0.005	25.78	1.08	0.003	28.73
	(1.17)		(0.004)	[0.06]	(0.74)	(0.003)	[0.03]
	[0.11]		[0.26]	{16}	[0.14]	[0.42]	{16}
RSMALL	3.61	−0.63	0.007	26.94			
RTSE2	(1.98)	(1.52)	(0.005)	[0.03]			
	[0.07]	[0.68]	[0.15]	{15}			
	1.88	$b \equiv 0$	0.005	25.92	1.04	0.003	29.13
	(1.19)		(0.004)	[0.06]	(0.75)	(0.003)	[0.02]
	[0.11]		[0.24]	{16}	[0.17]	[0.39]	{16}

— *continued*

TABLE 5 (*Continued*)

Panel B: Recursive utility model				
Assets	γ	ν	θ	J_T
RSMALL	1.01	0.79	−0.000	24.22
RET1	(0.92)	(0.20)	(0.003)	[0.06]
	[0.27]	[0.00]	[0.91]	{15}
RSMALL	0.96	0.80	−0.010	24.39
RET2	(0.89)	(0.19)	(0.002)	[0.06]
	[0.28]	[0.00]	[0.00]	{15}
RSMALL	1.00	0.79	−0.010	24.36
RTSE2	(0.90)	(0.20)	(0.003)	[0.06]
	[0.27]	[0.00]	[0.00]	{15}

Estimation of the following equation is based on Hansen's (1982) generalized method of moments,

$$E_t(m_{t+1}(R_{j,t+1} + \theta)|Z_t) = 1,$$

where Z_t contains a constant and two lags each of term premium, default premium, the nominal returns on the Section I of TSE stocks, and the nominal interest rate. The standard errors reported in parentheses are based on the simple covariance matrix estimator outlined in Hansen (1982). The p-value in brackets tests the hypothesis that the estimated parameter equals zero. The degree of freedom df (reported in curly brackets) is the number of moment conditions minus the number of parameters to be estimated. The J_T-statistic tests whether the over-identifying restrictions of the model are true with the degrees of freedom, df. The statistic, A_T, which is asymptotically $\chi^2(1)$-distributed tests the exclusion restriction, $b=0$. For each estimation, set $\beta = 0.99$. The test assets are as defined in Table 3. The reported estimation for $b \equiv 0$ implies the time-separable model.

where θ can be interpreted as a pricing error similar to Jensen's alpha. Using the same previous set of test assets and the set of conditioning information, the restrictions imposed by $\langle 12 \rangle$ are tested via GMM. Table 5 reports the estimates for $\{b, \gamma, \theta\}$. The Jensen's alpha for the non-separable and the Abel (1990) model are distinctively positive but are negative for the recursive utility model. The reported sign of the pricing error substantiates the assertions in Constantinides (1990) that non-separability enhances the ability of the stochastic discount factor to explain the time-variation in the equity premium mostly through its impact on the higher moments. The estimated θ suggests that the average equity premium that provides the best fit to the model in $\langle 12 \rangle$ is higher (lower) than the historical average equity premium for the non-separable (unexpected) utility model.

IV. Concluding remarks

This article provides an empirical investigation of a large class of dynamic asset pricing models. We have both examined and contrasted the empirical fit of five asset pricing models: (i) the standard expected utility model; (ii) Abel's (1990) model with consumption externality; (iii) the time non-separable model; (iv) the recursive utility model; and (v) the ad-hoc factor model. Using Japanese stock and bond market data, the asset pricing restrictions implied by each model were subjected to the Hansen–Jagannathan (1991) (and Cecchetti *et al.*, 1994) volatility bounds tests and the Hansen–Jagannathan (1994) specification error tests. The *J*-statistic goodness-of-fit criteria was applied to test the validity of each asset pricing model. The preference parameters governing each stochastic discount factor were estimated using generalized method of moments. These empirical exercises not only helps us to understand the aggregate risk-taking behavior of the Japanese investors but also how it is related to the intertemporal behavior of security returns and consumption growth.

The main empirical findings are as follows. First, on the basis of the Hansen–Jagannathan (1994) specification error tests we can conclude that: (a) the implied coefficient of relative risk aversion (from the expected utility model) is around 8.50. (b) The non-separability parameter consistent with the minimized specification error is about 0.55 and the concavity parameter is 2.50. (c) Japanese consumers prefer early resolution of uncertainty, and the intertemporal elasticity of substitution with respect to the real interest rate is small. (d) The specification error corresponding to the non-expected utility asset pricing model differs only slightly from the expected utility model and is much larger compared to the ad-hoc linear factor and the time non-separable asset pricing models. The data are thus rejecting the asset pricing model based upon the recursive preference structure. Asset pricing models based upon non-separability in consumption expenditures (i.e. habit formation) and the ad-hoc factor pricing model performed the best in terms of the minimized distance criteria.

Second, the results from the Hansen–Jagannathan (1991) volatility bounds tests are consistent with those from the specification error tests. For example, the minimum risk aversion parameter required to satisfy the Hansen–Jagannathan volatility bounds is 12.0. In the context of Japanese capital markets, this means that consumption growth varies too little to explain the equity premium unless the coefficient of relative risk aversion is high. The inability of asset pricing models to generate volatile stochastic discount factors is also applicable to Japanese security markets. Third, while testing the Euler equation implications implied by the non-separable asset pricing model, it was found that the empirical support for durability was weak in monthly data. The effect of durability was, however, dominated by habit forming preferences in quarterly consumption data. This set of results is similar to the findings of Ferson and Constantinides (1991) that the time-series properties of consumption and half-lives of durability and habit formation are related to the observation frequency. Furthermore, the magnitude and sign of the estimated Jensen's alpha is also supportive of the habit formation model. The overall conclusion is that stochastic discount factors that take into account habit forming preferences do better in explaining the empirically observed asset prices.

The search for better performing stochastic discount factors can be extended to asset pricing models that incorporate market frictions. This line of research

can appeal to the characterizations in Hansen *et al.* (1995) and He and Modest (1995), and is left to a follow-up empirical project.

Appendix: the data description

Reliable capital market data are available for Japan starting only in the 1970s, so the 1971:1–1990:12 sample period is included for our study. For the empirical work, define the following variables (see Table A1):

- c_t: per capita, seasonally adjusted real non-durable consumption. c_t^s is per capita seasonally unadjusted real nondurable consumption. $CGSA_t$ and $CGSN_t$ are, respectively, one plus the percentage changes in c_t and c_t^s. Source: *OECD Main Economic Indicators* (provided on a diskette by *ESTIMA*).
- $NINT_t$: short-term nominal interest rate. Through November 1980, it is the daily average of the overnight call rate. After that, the one month Gensaki rate is used. Source: Ibbotson and Hamao (1991).
- $TERM_t$: corresponds to the term premium. It is the excess return on an index of long-term government bonds having a maturity of 10 years over the short-term nominal interest rate. Source: Ibbotson and Hamao (1991).

TABLE A1. Summary statistics

Variable	Mean	SD	$\rho 1$	$\rho 2$	$\rho 3$	$\rho 6$	$\rho 12$	$\rho 24$	Q(24)	p-value
CGSN	1.0063	0.0891	−0.35	−0.28	0.15	−0.00	0.89	−0.03	569.74	0.00
CGSA	1.0018	0.0195	−0.37	−0.08	0.20	0.14	−0.04	−0.10	169.32	0.00
RTSE1	1.0089	0.0514	0.09	0.11	0.02	0.05	0.06	0.04	26.22	0.34
RTSE2	1.0067	0.0537	0.24	0.09	−0.04	0.03	0.07	0.10	38.58	0.03
RSMALL	1.0135	0.0567	0.18	0.07	0.04	0.02	0.10	0.05	32.45	0.12
RPREM	0.0073	0.0492	0.05	0.10	0.03	0.04	0.03	0.06	20.22	0.68
TERM	0.0007	0.0174	0.16	0.00	−0.05	−0.03	0.06	−0.00	41.40	0.02
DEF	0.0002	0.0103	−0.06	−0.01	−0.07	0.08	0.08	−0.09	34.31	0.08
NINT	1.0055	0.0019	0.97	0.95	0.92	0.79	0.42	0.10	1782.08	0.00
NTSE1	1.0128	0.0440	0.04	0.09	0.02	0.03	0.03	0.06	18.62	0.77
INF	0.0038	0.0103	0.20	0.04	−0.04	0.40	0.62	0.28	320.84	0.00
DYIELD	0.0165	0.008	0.97	0.94	0.91	0.83	0.67	0.58	3046.75	0.00

$CGSN_t$ is one plus the growth rate of per capita seasonally unadjusted real consumption. $CGSA_t$ is one plus the growth rate of per capita seasonally adjusted real consumption. $NINT_t$ measures the short-term nominal interest rate. $TERM_t$ corresponds to the term premium. It is the excess return on an index of government bonds having maturity of 10 years over the short-term nominal interest rate. DEF_t tracks the default premium, and is the return on an index of corporate bonds having a maturity of 10 years net of returns on an index of government bonds having a maturity of 10 years. $RTSE1_t$ is the real return on the value weighted index of stocks traded on Section I of the Tokyo Stock Exchange. $RTSE2_t$ corresponds to the real return on the value weighted index of Tokyo Stock Exchange Section II stocks. $RSMALL_t$ tracks the value weighted real return on the smallest quintile of stocks traded on Section I of the Tokyo Stock Exchange. $RPREM_t$ measures the market risk premium and is the excess return of $RTSE1$ over the short-term real interest rate. INF_t is the CPI inflation rate. $\rho(L)$ is the autocorrelation coefficient, and $Q(L)$ is the Ljung–Box statistic at lag L. The $Q(L)$-statistic is distributed $\chi^2(L)$ uder null hypothesis of no serial correlation. The sample period is 1971:1–1991:12 (240 observations).

- DEF_t: tracks the default premium, and is the excess return on an index of corporate bonds over an index of government bonds (both indexes have a maturity of 10 years). Source: Ibbotson and Hamao (1991).
- $RTSE1_t$: real return (dividend inclusive) on a value-weighted index of Tokyo Stock Exchange Section I stocks. The real returns are nominal returns minus consumer price index inflation. The consumer price index is used to deflate nominal returns since the consumption deflator is unavailable. This is unlikely to bias the empirical results as Japan has experienced moderate levels of inflation. Source: Ibbotson and Hamao (1991).
- $RTSE2_t$: corresponds to the value-weighted, dividend inclusive, real returns on the Tokyo Stock Exchange Section II stocks. Source: Ibbotson and Hamao (1991).
- $RSMALL_t$: value-weighted, dividend inclusive, real return on the smallest quintile of Tokyo Stock Exchange Section I stocks. The data source is again, Ibbotson and Hamao (1991).
- $RLTGB_t$: real return on the index of long-term government bonds having a maturity of 10 years. Source: Ibbotson and Hamao (1991).
- $RLTCB_t$: real return on the index of long-term corporate bonds having a maturity of 10 years. Source: Ibbotson and Hamao (1991).
- $RET1_t$: synthetic return constructed by investing a fraction in $RSMALL_t$ and the rest in $RTSE1_t$. The weight is lagged term premium multiplied by 10.
- $RET2_t$: synthetic return constructed by investing a fraction in $RSMALL_t$ and the rest in $RLTGB_t$. The weight is lagged default premium multiplied by 10.
- $DYIELD_t$: tracks the dividend-yield. It is the sum of the previous 12 months dividends divided by the time-t price of the value-weighted index of Tokyo Stock Exchange Section I stocks.

Notes

1. The period t utility function is: $U(c) = (1-\gamma)^{-1}\,(c_t/c^*_{t-1})^{1-\gamma}$, where c^*_t represents aggregate consumption.
2. The utility function in their empirical investigations is of the power form: $U(c) = (1-\gamma)^{-1}(c_t - bc_{t-1})^{1-\gamma}$.
3. For the Hagnsen–Jagannathan specification error and the Hansen–Jagannathan (1991) volatility bounds tests, we tried other combinations of test assets that included the rate of return on corporate bonds. But since including other test assets resulted in the same qualitatative conclusions, to save space, the results from other test assets are not reported.
4. The absolute value of Sharpe ratio, for any excess return r, given by

$$\frac{\sigma(m)}{E(m)} \geq \frac{E(r)}{\sigma_r}$$

puts a lower bound on the variability of the IMRS relative to its expected value.
5. Similar results for the recursive utility model are reported (using US data) in the volatility bounds test of Cochrane and Hansen (1992, Figure 3.2).
6. The standard errors were insensitive to the lag structure employed in the Newey–West correction procedure.
7. The implementation of the time non-separable model requires knowledge of the statistical process followed by per capita consumption growth. This is well approximated by an AR-1 process. For an alternate procedure, see Hansen *et al.* (1995). We thank Nelson Mark for providing the *GAUSS* code used in their study (Cecchetti *et al.*, 1994).
8. In implementing this test, we use the equivalent specification:

$$\delta = \left\{[\underline{1} - E(m\underline{R})]'E(\underline{R}\underline{R}')^{-1}[\underline{1} - E(m\underline{R})]\right\}^{1/2},$$

which has the advantage of not requiring the data on the price vector $\underline{q}$. Here $\underline{R}$ denotes the rate of return vector for the test assets.

9. Other variables such as dividend-yield and yield-based measures were also included in the set of conditioning information in our pretrial estimations but that resulted in similar parameter estimates. So to save space, exclusive attention is devoted to the above set of conditioning instruments.
10. In this model, ε_t is a function of $R_{j,t}$, c_t, and c_{t-1} and $\varepsilon_t \sim$ MA(1). Consequently, the weighting matrix is adjusted to account for the moving average term, as in Hansen (1982). To avoid degenerate solutions ε_t is divided by $1 - b\beta$.

References

Abel, A. B. (1990) Asset prices under habit formation and catching up with the Joneses. *American Economic Review* **80**, 38–42.

Bakshi, G. S. and Chen, Z. (1996) The spirit of capitalism and stock market prices. *American Economic Review* **86**, 133–157.

Bakshi, G. S., Chen, Z. and Naka, A. (1995) Production-based asset pricing in Japan. *Pacific-Basin Finance Journal* 3, 217–240.

Braun, P. A., Constantinides, G. M. and Ferson, W. E. (1993) Time non-separability in aggregate consumption: International evidence. *European Economic Review* **37**, 897–920.

Campbell, J. Y., and Hamao, Y. (1992) Predictable stock returns in the US and Japan: A study of long term capital market integration. *Journal of Finance* **47**, 43–69.

Cecchetti, S. G., Lam, P.-S. and Mark, N. C. (1994) Testing volatility restrictions on intertemporal marginal rates of substitution implied by euler equations and asset returns. *Journal of Finance* **49**, 123–152.

Cochrane, J. H. (1994) Discrete-time empirical finance, Working Paper. University of Chicago.

Cochrane, J. H. and Hansen, L. P. (1992) Asset pricing explorations for macroeconomics. *NBER Macroeconomics Annual*, pp. 115–16.5. MIT Press, Cambridge, MA.

Constantinides, G. M. (1990) Habit formation: a resolution of the equity premium puzzle. *Journal of Political Economy* **98**, 519–543.

Dunn, K. B. and Singleton, K. J. (1986) Modelling the term structure of interest rates under non-separability of preferences and durability of goods. *Journal of Financial Economics* **17**, 27–55.

Eichenbaum, M. and Hansen, L. P. (1990) Estimating models with intertemporal substitution using aggregate time-series data. *Journal of Business and Economic Statistics* **8**, 53–69.

Epstein, L. G. and Zin, S. E. (1989) Substitution, risk aversion, and the temporal behavior of consumption and asset returns: a theoretical framework. *Econometrica* **57**, 937–969.

Epstein, L. G. and Zin, S. E. (1991) Substitution, risk aversion, and the temporal behavior of consumption and asset returns: an empirical analysis. *Journal of Political Economy* **99**, 263–286.

Ferson, W. E. (1994) Theory and empirical testing of asset pricingmodels. Ch. 6. In *The Finance Handbook*, ed R. Jarrow, W. Ziemba and V. Maksimovic, 1994.

Ferson, W. E. and Constantinides, G. M. (1991) Habit persistence and durability in aggregate consumption: empirical tests. *Journal of Financial Economics* **29**, 199–240.

Ferson, W. E. and Foerster, S. R. (1994) Finite sample properties of the generalized method of moments in tests of conditional asset Pricing models. *Journal of Financial Economics* **36**, 29–55.

Ferson, W. E. and Harvey, C. R. (1992) Seasonality and consumption-based asset pricing. *Journal of Finance* **47**, 511–552.

Hansen, L. P. (1982) Large sample properties of generalized method of moments estimators. *Econometrica* **45**, 1029–1079.

Hansen, L. P., Heaton, J. and Luttmer, E. G. (1995) Econometric evaluation of asset pricing models. *Review of Financial Studies* **8**, 237–274.

Hansen, L. P., Heaton, J. and Yaron, A. (1994) Finite sample properties of some alternative GMM estimators, Working Paper. University of Chicago.

Hansen, L. P. and Jagannathan, R. (1991) Implications of security market data for models of dynamic economies. *Journal of Political Economy* **99**, 225–262.

Hansen, L. P. and Jagannathan, R. (1994) Assessing specification errors in stochastic discount factor models. Federal Reserve Bank of Minneapolis, Research Department Staff Report No. 167.

Hansen, L. P. and Richard, S. F. (1987) The role of conditioning information in deducing testable restrictions implied by dynamic asset pricing models. *Econometrica* **55**, 587–613.

Hansen, L. P. and Singleton, K. J.(1982) Generalized instrumental variable estimation of non-linear rational expectations models. *Econometrica* **50**, 1269–1286.

He, H., and Modest, D. M. (1995) Market frictions and consumption-based asset pricing. *Journal of Political Economy* **103**, 94–117.

Heaton, J. (1995) An empirical investigation of asset pricing with temporally dependent preferences. *Econometrica* **63**, 681–717.

Heaton, J., and Lucas, D. J. (1992) The effects of incomplete insurance markets and trading costs in a consumption-based asset pricing model. *Journal of Economic Dynamics and Control* **16**, 601–620.

Ibbotson, R. and Hamao, Y. (1991) *Stocks, Bonds and Inflation Japan Year Book*. Ibbotson Associates Inc.

Jagannathan, R. and Wang, Z. (1996) The CAPM is alive and well. *Journal of Finance* **51**, 3–53.

Newey, W. K. and West, K. D. (1987a) A simple positive semi-definite heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica* **55**, 703–708.

Newey, W. K. and West, K. D. (1987b) Hypothesis testing with efficient method of moments estimation. *International Economic Review* **28**, 777–787.

Porter, R. and Wheatley, S. (1995) Time-non-separable preferences or artifact of temporal aggregation?. Working Paper, University of Chicago and University of South Wales, 1995.

Ross, S. A., (1978) A simple approach to the valuation of risky streams. *Journal of Business* **51**, 415–477.