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Intertemporal Substitution and Consumer Durables: An Analysis Based on Japanese Data *

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Abstract

Several studies using Japanese aggregate data suggest that the intertemporal elasticity of substitution for Japanese consumers is far larger than that for U.S. consumers. The available estimates for Japan, however, do not allow for both intratemporal substitution between non-durables and durables and time aggregation. After corrections for removing these two effects are made, this paper shows that the point estimates for Japan are still larger than those for the U.S., but are unlikely to be much above one. This result suggests that Japan's high saving rate may be explained by the magnitude of the intertemporal elasticity of substitution.

Key words: Intertemporal Elasticity of Substitution, Consumer Durables, Japanese Saving Behavior, Cointegration, Generalized Method of Moments

JEL classification: C22, E21

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1 Introduction

The intertemporal elasticity of substitution (IES) is a key parameter that characterizes household consumption and saving behavior, and plays a crucial role in policy and welfare assessments. For example, the sensitivity of household savings to real interest rates is weak if the IES is small and the effects of fiscal policies are crucially different depending on the magnitude of the IES. However, in the empirical literature involving U.S. data, since the IES was frequently estimated to be very small or negative (see, e.g., Hall, 1988), the question as to why such results are obtained or how the magnitude of the estimate is valid has been examined from various aspects (see, e.g., Hansen and Singleton, 1996; Beaudry and van Wincoop, 1996; Stock and Wright, 2000; and Neely, Roy, and Whiteman, 2001).

While most work has not theoretically predicted the direction of the bias in estimates of the IES, some recent studies find positive and significant estimates of the IES by allowing for possible factors that cause a downward bias. One theoretical reason for explaining it is the existence of the intratemporal substitutability between non-durables and durables. If such a relationship exists, by using the model that ignores the role of durables, the IES will be underestimated, because an increase in real interest rates decreases the growth rate of non-durables when it increases the user cost of durables. Ogaki and Reinhart (1998) focus on this misspecification bias, and find estimates of the IES around 0.3–0.5, using U.S. aggregate data. Another theoretical explanation is given by Vissing-Jørgensen (2002), who argues that when the consumption growth rate of non-asset holders is not positively correlated with the predictable change in asset return, the IES, based on the consumption growth rate of aggregate data (i.e., all households), will be underestimated. Interestingly, in her paper, Vissing-Jørgensen finds that IES estimates for asset holders are around 0.3–1.0, while those for non-asset holders are small and not significantly different from zero, using micro data from the U.S. Consumer Expenditure Survey (CEX).

These findings suggest that allowing for two different aspects of aggregation, namely, non-separability across goods and heterogeneity among consumers, is equally important for obtaining reasonable IES estimates. However, this implication is based only on U.S. data. It is thus of interest to see if such misspecification biases in estimates of the IES are confirmed even in other data sets.

In this paper, we use Japanese aggregate data to reevaluate the importance of allowing for non-separability across goods in estimating the IES. We examine whether the IES is estimated to be significantly greater than zero, and if so, how the value is reasonable for Japanese consumers. This attempt is important because, while so far some empirical findings from Japanese data has helped deepen our understanding about consumption and saving decisions, little attention has been given to the magnitude of the IES for Japanese consumers.[1] Given careful and detailed research on the contrasts between the U.S. and Japanese saving behaviors (see, e.g., Hayashi, 1997), one possible conjecture for the above question is that the IES for Japanese consumers is probably larger than that for U.S. consumers, because consumers will save more proportionally when the IES is large. To make the empirical results between the U.S. and Japan comparable, we adopt the same method as Ogaki and Reinhart (1998), and

[1] In contrast to the case of the U.S., most previous studies have tended to place the major focus on tests of overidentifying restrictions or model diagnoses using the methods of Hansen and Jagannathan (1991, 1997), rather than on the estimation of the IES itself (see, e.g., Saito, 1999, and Hori, 2000 for comprehensive surveys). Although a few empirical studies, e.g., Kitamura and Fujiki (1997) and Nakano and Saito (1998), find estimates of the IES around 0.4, given a serious lack of empirical investigation, it seems fair to state that there is no consensus on values of the IES for Japanese consumers among researchers.

investigate an empirical validity of this conjecture.

A further advantage of using their method is that we can analyze explicitly the cause of an extremely large IES reported in Hamori (1992, 1996).[2] In particular, we focus on the fact that such previous studies define utility on total consumption in estimating the Euler equation, thereby examining whether the large IES can be explained as the result of imposing the untenable assumption about aggregation across goods.[3]

In this paper, we find estimates of the IES around 0.7–1.0 for Japanese consumers by allowing for the intratemporal substitution between non-durables and durables. We find, however, that ignoring the intratemporal substitution yields very similar estimates to the U.S. case, as in Kitamura and Fujiki (1997) and Nakano and Saito (1998), and that the misspecification of aggregation across goods yields large estimates, as in Hamori (1992, 1996). Thus, the analysis of the present paper presents further evidence in favor of the claim of Ogaki and Reinhart (1998) that ignoring the intratemporal substitution leads to a downward bias in estimates of the IES; in addition, it is partly successful in reconciling our findings of the IES with those of previous Japanese studies.

However, as in Hall (1988) and Vissing-Jørgensen (2002), we interpret the IES as the elasticity of the consumption ratio to the real interest rate. Therefore, while our findings can be viewed as guides to reasonable estimates of the IES for Japanese consumers, they should not be taken as direct evidence for settling the debate over equity premium puzzles.

The paper is organized as follows. Following Ogaki and Reinhart (1998), Section 2 presents

[2] In his papers, Hamori typically reports IES estimates that far exceed 10.

[3] As the other contributing factor, Saito (1998) argues that market incompleteness could help explain the puzzling finding of a large IES in Japan.

the model and gives an outline of a two-step approach that combines a cointegrating regression with generalized method of moments (GMM). Section 3 discusses characteristics of the data used in this paper. Section 4 reports estimation results and discusses their interpretations. Section 5 concludes and provides some suggestions for future research. Finally, additional tests for ensuring our results are given in the Appendix.

2 The Model

This section briefly reviews Ogaki and Reinhart's (1998) model with nonseparable preferences in non-durables and durables, and explains the two-step approach for estimating preference parameters. Suppose that a representative consumer maximizes the following lifetime expected utility at time 0:

$$U = E_0 \left[\sum_{t=0}^{\infty} \beta^t \left(\frac{\sigma}{\sigma - 1} \right) \left(u(t)^{1-\frac{1}{\sigma}} - 1 \right) \right], \tag{1}$$

where β is the subjective discount factor, $E_t[\cdot]$ is the expectations operator conditional on the information available at time t, $\sigma > 0$ is the intertemporal elasticity of substitution (IES), and $u(t)$ is the period utility function at time t. To allow for the intratemporal substitution effect, we assume that the consumer drives utility in each period according to the following constant elasticity of substitution (CES) function:

$$u(t) = \left(aC_{1t}^{1-\frac{1}{\epsilon}} + S_t^{1-\frac{1}{\epsilon}} \right)^{\frac{1}{1-\frac{1}{\epsilon}}}, \tag{2}$$

where C_{1t} is the consumption of non-durables, S_t is the service flow from the purchases of durables, $\epsilon > 0$ is the elasticity of substitution between C_{1t} and S_t, and $a > 0$ is some number that denotes the weight attached to non-durables in the period utility function. The relation between the purchases of durables and the service flow is given by

$$S_t = C_{2t} + \delta C_{2t-1} + \delta^2 C_{2t-2} + \cdots, \tag{3}$$

where C_{2t} is the real consumption expenditure for durables at time t, and $1-\delta$ is the depreciation rate of durables.

Let R_{t+1} be the gross return on any asset between t and $t+1$, and let P_t be the purchase price of durables in terms of non-durables. Then the Euler equation with respect to C_{1t} is

$$E_t\left[\beta\left(\frac{mu_{t+1}}{mu_t}\right)R_{t+1}\right] = 1, \tag{4}$$

where

$$mu_t = aC_{1t}^{-\frac{1}{\epsilon}}\left(aC_{1t}^{1-\frac{1}{\epsilon}} + S_t^{1-\frac{1}{\epsilon}}\right)^{\frac{\sigma-\epsilon}{\sigma(\epsilon-1)}}. \tag{5}$$

For an optimum, moreover, the following two first-order conditions are necessary:

$$Q_t = a^{-1}\left(\frac{S_t}{C_{1t}}\right)^{-\frac{1}{\epsilon}}, \tag{6}$$

$$P_t = \frac{E_t\left[\sum_{\tau=0}^{\infty}\beta^{\tau}\delta^{\tau}mu_{2t+\tau}\right]}{mu_t}, \tag{7}$$

where Q_t is the user cost for the service flow of durables, given by

$$Q_t = P_t - \delta E_t\left[\frac{\beta P_{t+1} mu_{t+1}}{mu_t}\right], \tag{8}$$

and the marginal utility of S_t is expressed as

$$mu_{2t} = S_t^{-\frac{1}{\epsilon}}\left(aC_{1t}^{1-\frac{1}{\epsilon}} + S_t^{1-\frac{1}{\epsilon}}\right)^{\frac{\sigma-\epsilon}{\sigma(\epsilon-1)}}. \tag{9}$$

Equation (6) is the condition that the user cost is equal to the marginal rate of substitution between the service flow and non-durables. More particularly, it is used, together with equation (8), to calculate the parameter, a.[4] On the other hand, equation (7) is the condition that the relative price, P_t, is equal to the marginal rate of substitution between non-durables and durables. As described below, this condition is used to derive the restriction that implies cointegration. Finally, in order to specify the supply side of the model in the simplest way, consider an endowment economy without production. Then, market-clearing conditions for non-durables and durables are $C_{1t} = C_{1t}^*$ and $C_{2t} = C_{2t}^*$, where C_{1t}^* and C_{2t}^* denote the endowments of non-durables and durables, respectively.

On the basis of the model outlined, Ogaki and Reinhart's two-step estimation procedure can be summarized as follows. Consider the case in which both $\ln C_{1t}$ and $\ln C_{2t}$ are difference stationary in equilibrium. Multiplying both sides of equation (7) by $(C_{1t}/C_{2t})^{-1/\epsilon}$, we obtain

$$P_t \left(\frac{C_{1t}}{C_{2t}}\right)^{-\frac{1}{\epsilon}} = E_t \left[\left(\frac{1}{a}\right) \sum_{\tau=0}^{\infty} \beta^{\tau} \delta^{\tau} \left(\frac{S_{t+\tau}}{C_{2t}}\right)^{-\frac{1}{\epsilon}} \left(\frac{C_{1t}}{C_{1t+\tau}}\right)^{-\frac{1}{\epsilon}} \left(\frac{mu_{t+\tau}}{mu_t}\right) \right], \tag{10}$$

where the expression on the right-hand side follows from equations (3), (5), and (9). If the discounted sum on the right-hand side is stationary, and if the random variables (in the consumer's information set) used to form the conditional expectation on the right-hand side are stationary, then $P_t(C_{1t}/C_{2t})^{-1/\epsilon}$ is stationary since the right-hand side is stationary. [5] After taking the natural logarithm of both sides of equation (10) and rearranging, we can use the result that $\ln(C_{1t}/C_{2t}) - \epsilon \ln P_t$ is stationary as the restriction that summarizes the information from the demand side of the model. We shall refer to this restriction as the

[4] In this paper, the parameter a is not of immediate interest. In addition, since equation (8) involves the conditional expectation operator, it requires any approximation to calculate the user cost. For these reasons, we focus on another first-order condition to derive a restriction underlying our empirical specification.

[5] In equation (10), the stationarity of the first two variables, $S_{t+\tau}/C_{2t}$ and $C_{1t}/C_{1t+\tau}$, follows directly from the difference stationarity of $\ln C_{1t}$ and $\ln C_{2t}$, while the stationarity of $mu_{t+\tau}/mu_t$ is not necessarily implied by the model. As Ogaki and Reinhart illustrate, however, it seems safe to state that the stationary assumption of the growth rate of marginal utility is valid at least empirically. In fact, as in the Appendix, we found strong evidence that supports this assumption for the growth rate of estimated marginal utility.

stationary restriction. If the variable $\ln(C_{1t}/C_{2t})$ is difference stationary, it follows from this stationary restriction that $\ln P_t$ is also difference stationary. As a result, we obtain a cointegration relationship between $\ln(C_{1t}/C_{2t})$ and $\ln P_t$.

The two-step procedure exploits this cointegration relationship to estimate the intratemporal elasticity of substitution, ϵ, in the first step. Specifically, we consider the following cointegrating regression:

$$\ln(C_{1t}/C_{2t}) = \theta + \epsilon \ln P_t + u_t, \tag{11}$$

where θ is a constant term, and u_t is a stationary variable with zero mean. It can be easily shown that appropriate estimation of this cointegrating regression yields a super-consistent estimate of ϵ. We can, therefore, estimate intertemporal parameters, σ and β, by applying Hansen's (1982) GMM to the Euler equation (4) into which the estimate of ϵ from regression (11) was plugged. To turn to this step of the procedure, the disturbance is defined as

$$v_{t+1} = \beta \left(\frac{\widehat{mu}_{t+1}}{\widehat{mu}_t} \right) R_{t+1} - 1 \tag{12}$$

such that $E_t[v_{t+1}] = 0$, where '^' denotes that the value of ϵ is fixed at the estimate from regression (11). Letting Z_t denote a matrix of variables that are in the information set at time t, we obtain the following moment condition from the Euler equation (4):

$$E[Z_t v_{t+1}] = 0. \tag{13}$$

In the second step, this moment condition is exploited for estimating the intertemporal parameters and evaluating the model, as in the standard GMM approach.

3 Data and Pretests

We use quarterly data from the Annual Report on the National Accounts, compiled by the Economic and Social Institute of Cabinet Office (formerly the Economic Planning Agency) of the Japanese government. The variables C_{1t} and C_{2t} in the model are taken as the real per capita expenditures on non-durables and durables, respectively. To obtain the per capita series, real consumption is divided by total population (averaged over each quarter) in the Monthly Report on Current Population Estimates, compiled by the Statistical Bureau of the Management and Coordination Agency of the Japanese government. The relative price, P_t, is defined using the implicit deflators for each good. All series are seasonally adjusted by the census X-11 method.[6] The remaining variables are not available directly from the Reports; therefore, they are constructed in the following ways.[7]

The service flow series, S_t, is calculated from equation (3) using data on the real per capita expenditure on durables, the stock of durables at the beginning of the year 1971 (as the initial value of S_t), and $\delta = 0.94$. Here the assumption of $\delta = 0.94$ for the quarterly data means depreciation at the annual rate of about 22 percent. This choice of δ is the same as that of Ogaki and Reinhart (1998) using the U.S. data, and is also roughly consistent with the result of Hayashi (1997, chap. 11) calculated using the Japanese data.[8]

For the user cost series, Q_t, a simple approximation is used after substituting equation (4) into equation (8). That is, by assuming the independency between the expected relative

[6] This is because seasonally adjusted data on non-durables and durables is not published.

[7] The process of data construction follows that of Ogaki and Reinhart (1998). Quarterly data in the National Accounts are available from the second quarter of 1970; therefore, the data of the period starting 1971 are employed in the construction.

[8] Similar choices are also made in Mankiw (1985) and Fauvel and Samson (1991). These authors assume depreciation rates of 20 percent for the U.S. data and 25 percent for the Canadian data, respectively.

price and the inverse of the expected gross return, we utilize

$$Q_t \cong P_t - \frac{\delta E_t[P_{t+1}]}{E_t[R_{t+1}]}. \tag{14}$$

Here, to calculate the gross return, the call rate is used as nominal interest rates, which are converted into real rates by the implicit deflator of non-durables. The expected variables on the right-hand side are constructed by estimating a bivariate vector autoregression (VAR) model with three lags for the realized real interest rate and the growth rate of the relative price.[9] With these variables in hand, the parameter, a, is now easy to calculate:

$$a = \exp\left[\frac{\ln(C_{1t}/S_t)}{\epsilon} - \ln Q_t\right]. \tag{15}$$

Using this formula derived from equation (6), the value of a is calculated as the exponential of the sample mean of the variable in the square bracket.

Table 1 reports standard deviations and correlation coefficients for the constructed series. Standard errors for these estimates are calculated by applying GMM with a VAR(1) prewhitening technique of Andrews and Monahan (1992) (see Section 8.4 of Ogaki, 1993 for details). The sample period is from the first quarter of 1975 to the fourth quarter of 1997.

The following can be seen from each panel of Table 1. First, the standard deviation of the growth rate of durables is greater than that of non-durables, while the standard deviations of non-durables and the service flow from durables are fairly similar. However, despite such a similarity in terms of the standard deviation, the last row of Panel (A) suggests that Hicksian aggregation across goods does not hold. That is, the user cost relative to the price of non-durables, which corresponds to the relative price of the service flow and non-durables,

[9] Results reported below are not sensitive to the choice of the lag order in the VAR model.

is not constant over time. This point is especially important because, as mentioned in the Introduction, most of the previous studies using the Japanese data estimate the IES based on the single-good model for total consumption (defined as the sum of non-durables plus services, durables, and semi-durables). In other words, they ignore not only the intratemporal substitution between different goods but also separability across goods. This misspecification may explain why the earlier studies have found an extremely large IES. In the next section, we will show indeed that aggregation across goods gives a crucial bias in estimates of the IES.

Second, the correlation between the growth rate of the user cost, $\ln Q_t - \ln Q_{t-1}$, and the expected real interest rate is statistically positive at a high significance level. This result is consistent with Ogaki and Reinhart (1998), suggesting that there may be a substantial downward bias in estimates of the IES even when we use the Japanese data. To make this result more precise, the last row of Panel (B) reports the correlation coefficient between the expected relative price and the inverse of the expected gross return. Since this coefficient is not significantly different from zero, we expect that the approximation of equation (14) does not cause a serious problem in practice.

To sum up, Table 1 suggests that, without incorporating the intratemporal substitution into the model, it is difficult to recover accurate point estimates of the IES from the Japanese data.

We close this section by examining the possibility of nonstationarity of the two variables, $\ln(C_{1t}/C_{2t})$ and $\ln P_t$. In this paper we use both the Augmented Dickey-Fuller (ADF) test developed by Said and Dickey (1984) and the J test proposed by Park (1990). These tests are designed so as to take trend stationarity as the alternative hypothesis, since both $\ln(C_{1t}/C_{2t})$ and $\ln P_t$ exhibit downward trends over the sample period. The test results are reported in Table 2. For both variables, we find that the null hypothesis of difference stationarity with drift cannot be rejected even at the 10% level. Hence, if the model is valid, we should find the cointegration relationship between these variables, as described in Section 2.

4 Estimation Results

4.1 Substitution between Non-durables and Durables

Given that both $\ln(C_{1t}/C_{2t})$ and $\ln P_t$ follow difference stationary processes with drift, the stationary restriction implies that these variables are deterministically cointegrated with a cointegrating vector $(1, -\epsilon)'$ in the terminology of Ogaki and Park (1998) and Campbell and Perron (1991). In other words, if the model is empirically valid, it implies that the two variables are stochastically cointegrated with the cointegrating vector, and that the deterministic cointegration restriction is satisfied in equilibrium. Therefore, we estimate equation (11) using Park's (1992) canonical cointegrating regression (CCR) procedure,[10] and apply Park's (1990) H(p,q) tests for the null hypotheses of stochastic cointegration and the deterministic cointegration restriction.

Table 3 reports estimation results of equation (11) and results of the H(0,1) and H(1,q) tests.[11] We consider here the regression with and without a dummy variable for the period of 1987:1 to 1993:4, before and after the collapse of the bubble economy. The point estimates

[10] Other asymptotically efficient estimators are proposed by Phillips and Hansen (1990), Saikkonen (1991), and Stock and Watson (1993). Since these all have the same limiting properties as the CCR estimator, the following estimation results are unlikely to depend on the choice of such estimation methods.

[11] We used the VAR(1) prewhitening technique of Andrews and Monahan (1992) to estimate the long-run variance matrix of the disturbances in the system. We report the third stage CCR estimates and the fourth stage H(p,q) statistics. For details of these points, see Park and Ogaki (1991) and Han (1996).

of ϵ are highly significant with the theoretically expected sign, indicating that the elasticity of substitution ranges from 2.208 to 2.537. Ogaki and Reinhart (1998) examine equation (11) using the U.S. quarterly data and report estimates of ϵ of about 1.2. The range of estimates for ϵ in Table 3 is therefore relatively large. This finding suggests that expenditures on non-durables and durables are more responsive to a change in the relative price than those in the US.

For both cases, neither the H(0,1) nor H(1,q) tests show strong evidence against the implications of cointegration mentioned above. Therefore, overall the CES specification of preferences is supported by the data. Because the dummy variable is significant at the 1% level, and because the H(p,q) test statistics are more favorable for the case with the dummy variable in terms of p-values, we adopt the estimate of ϵ obtained from the regression with the dummy variable in the following analysis.

4.2 The Intertemporal Elasticity of Substitution

Let us now turn to GMM estimation of σ. In the second step of our procedure, we impose the estimated values of ϵ and a upon the Euler equation (4) and apply GMM to the resulting Euler equation in order to obtain estimates of the IES. The instrumental variables used throughout all estimation procedures are a constant, the realized real interest rate, the growth rate of C_{1t}, and the growth rate of C_{1t}/C_{2t}.[12] To control for the time aggregation problem by the use of quarterly data, all instrument variables are lagged two periods, and the estimation of the optimal weighting matrix is conducted under the assumption that the disturbance follows an MA(1) process with an unknown coefficient (see Hall, 1988, and Hansen and Singleton,

[12] The following GMM results appear robust to other combinations of instrumental variables that include the growth rate of C_{2t} in place of the real consumption ratio.

1996 for details).

Table 4 reports the GMM results for the two-good model.[13] Following Ogaki and Reinhart (1998), we present estimation results when the discount factor, β, is fixed. This approach makes it possible to directly compare the results between the two-good and the one-good models.[14] We use here $\beta = 0.999$, 0.998, and 0.997, implying real interest rates of about 0.4%, 0.8%, and 1%, respectively, in an economy without growth.[15]

The first three rows of Table 4 report the results when β is fixed at the three values. For these base runs, we find that the Hansen's J-tests do not reject the overidentifying restrictions implied by the model. We also find that σ is estimated to be positive and significantly different from zero. These results provide strong support for our two-good model. One important feature to be emphasized here is that the standard error of σ becomes larger as we make the value of β smaller. Thus, even if much smaller values of β are used, a confidence interval for the estimate of σ does not eliminate the possibility of taking values of σ that are less than one.[16]

The fourth and fifth rows of Table 4 report the results of sensitivity analysis with respect

[13] The initial weighting matrix is an identity and the GMM estimation was iterated five times. We also imposed a penalty to eliminate exceptionally large values of σ. See footnote 9 of Ogaki and Reinhart (1998) for details. However, this penalty did not affect our estimation results.

[14] In addition, as Table 4 shows, β and σ are negatively correlated. For this reason, unless β is fixed, it is difficult to discuss the direction of changes in estimates of σ as in the analysis below.

[15] The sample mean of real interest rates employed in this paper is about 0.7%. This value seems similar to that in Kitamura and Fujiki (1997) and, in particular, Baba (2000) (it is 0.3% in Kitamura and Fujiki, and 0.8% in Baba), although their sample period is slightly different from ours. On the other hand, as Nakano and Saito (1998) discuss, when we use the data from the (Japanese) National Accounts, the estimated β often seem to become greater than one. Therefore, our choice of β may be appropriate at least for the present study based on the Japanese data. In fact, as in the U.S. case, choosing $\beta = 0.990$ (i.e., real interest rates of about 4%) led to the rejection of the model.

[16] However, when $\beta = 0.995$, we found that σ is estimated to be 1.645 with the standard error of 0.567. That is, the separability assumption ($\sigma = \epsilon$) may not be rejected for such a choice of β, because the point estimate of ϵ is in a range of one standard error of σ. In addition to points mentioned in footnote 15, this is another reason why β was fixed at a range of 0.997 to 0.999 in the present work.

to possible values of a and ϵ.[17] An increase in the value of a corresponds to attaching greater weight to non-durables in the period utility; a decrease in the value of ϵ corresponds to making the intratemporal substitution effect less important. In other words, these experiments both mean that we play down the role of durables in the estimation of the IES. Hence, we expect that such changes in a and ϵ yield smaller point estimates of the IES.

The fourth row of Table 4 reports the results when the value of a is increased,[18] and the fifth row of Table 4 reports the results when the value of ϵ is decreased by two standard errors, where the standard error is taken from Table 3. The results indicate that the above predictions are true for our point estimates of σ, revealing that it is important to allow for the role of durables to obtain more accurate estimates of the IES. This is consistent with Ogaki and Reinhart's findings.

Next, we examine how the ignorance of the intratemporal substitution can affect the estimates of the IES. Table 5 reports the GMM results for the one-good model. Panel (A) of Table 5 reports the results for the one-good model that assumes both $\sigma = \epsilon$ and $a = 1$ in the two-good model. In all cases, we find that the separability assumption ($\sigma = \epsilon$) yields smaller point estimates of σ. Again, it should be noted that when the value of β is small, the standard error of σ is large. These results can be viewed as a strong confirmation of the downward bias in the IES.

Panels (B) and (C) of Table 5 report the results for the one-good model in which ag-

[17] We also tried some different values of δ. As long as the values for δ were in a reasonable range, the results were very similar to those presented here. For example, when $\beta = 0.999$, we found an estimate of the IES of 0.767 for $\delta = 0.92$ and of 0.779 for $\delta = 0.96$.

[18] Here, after estimating the expected value for a and its standard error using the GMM with the VAR(1) prewhitening technique, we consider the case in which the GMM estimate of a is increased by two standard errors. The estimate of a is 7.644, and the standard error is 0.138.

gregation across goods is allowed. Our aim in this experiment is to examine what happens when we ignore even separability across goods in the estimation of the IES, which is a typical case in the empirical literature using Japanese data. If we can replicate the result of a large IES as in the literature, then we can provide one explanation that aggregation across goods is introducing a bias in the estimate of σ. Panel (B) reports the results when the sum of non-durables and the service flow from durable good purchases is taken as a single good, and Panel (C) reports the results when total consumption is taken as a single good. Given our results from Table 1 and Table 4, it is expected that the latter misspecification is more severe than the former one. The results indeed indicate that these misspecifications yield larger point estimates of σ, and that the point estimates of σ in Panel (C) have a stronger bias than those of σ in Panel (B). In addition, the J-test statistics strongly reject such one-good models.

4.3 Discussion

This subsection discusses how the above results are related to the literature. We focus on the magnitude of the estimated IES from the following three aspects. First, we have argued that the earlier finding of an extremely large IES may be attributable to the assumption of aggregation across goods. In the literature, this interpretation is not necessarily surprising. For example, Mankiw (1985) shows that an estimate of the IES for durables is larger than that of the IES for non-durables. An implication of Mankiw's result is that the use of a composite good formed from non-durables and durables is likely to yield larger point estimates of σ. Thus, the results in Panels (B) and (C) of Table 5 are consistent with this implication.

Second, we see that the magnitude of the IES under the separability assumption is in line with several earlier studies such as Kitamura and Fujiki (1997) and Nakano and Saito (1998). In particular, the results found in Panel (A) of Table 5 are very similar to those found in Kitamura and Fujiki (1997), who claim that point estimates of the relative risk aversion (RRA) parameter are around 1.6–3.7 (i.e., IES=0.3–0.6). The findings of the present paper, however, suggest that such an earlier result may be attributable to the downward bias under the separability assumption.

Third, after the corrections for the intratemporal substitution and time aggregation, we have found that the point estimates of the IES are around 0.7–1.0. This magnitude is larger than that in Ogaki and Reinhart (1998), but may be reasonable. In chapter 10 of Hayashi (1997), it is pointed out that Japan's national saving rate is not as high as commonly thought, but it is still higher than the U.S. rate, when calculated using the definition of the U.S. National Income and Product Accounts. If we note that larger values of the IES imply a greater willingness of a consumer to substitute consumption over time, we can interpret our result as evidence that confirms the finding of Japan's high saving rate in terms of measuring intertemporal substitution.

In addition, this difference in the magnitude of the IES between the U.S. and Japan may help to understand the growth of the postwar Japanese economy. According to the standard neoclassical growth model, the case of a large IES implies that consumers choose to consume relatively little when the capital stock is low; and an economy grows more rapidly due to the resulting high investment rate. In this respect, our finding may be useful to explain the rapid growth of Japan after the war.

5 Conclusion

In this paper, we have estimated the IES using Japanese aggregate data, and examined the question of why the value of the IES reported in the Japanese literature is far larger than the IES in the U.S. case. The aim of this paper is to offer additional insight with respect to the value of the IES through these attempts. We have found strong evidence that supports the claim of Ogaki and Reinhart (1998) that allowing for the intratemporal substitution between non-durable and durables is important in estimating the IES. In particular, our empirical results indicate that the IES is significantly above zero, and the point estimates of the IES are around 0.7–1.0. In this paper, we have argued that this finding that the IES for Japanese consumers is larger than that for U.S. consumers may be bound up with some remarkable features of postwar Japanese economy such as the high saving rate and rapid growth.

These results, however, still do not explain why there is the difference in the magnitude of the IES between the U.S. and Japan. To better understand the Japanese saving behavior, along the line of recent empirical research (e.g., Atkeson and Ogaki, 1996; Brav, Constantinides, and Geczy, 2002; Ogaki and Zhang, 2001; and Vissing-Jørgensen, 2002), allowing for the possibility of the wealth-varying IES and decreasing RRA may be one reasonable direction of extending our analysis. Furthermore, comparing the magnitude of the IES between high-income households and low-income households using consumption data of income quintile groups may be another interesting direction for future research. These extensions in turn may help explain the difference in the saving behavior of the two countries.

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Table 1
Standard Deviations and Correlations

Variable	Estimate	Standard error
(A) Standard deviation of growth rate		
Durables	0.0463	0.0096
Nondurables	0.0124	0.0020
Service flow from durables	0.0081	0.0025
User cost/Nondurables price	0.1198	0.0149
(B) Correlation		
User cost and Real interest rate	0.1993	0.0645
Relative price and Inverse of gross real return	0.2846	0.2647

[a] The service flow and the user cost of durables are calculated using $\delta = 0.94$.

Table 2
Unit Root Test Results

Variable	ADF(1)	ADF(4)	ADF(7)	J(1,5)
$\ln(C_{1t}/C_{2t})$	-1.390	-1.446	-1.619	2.175
$\ln P_t$	-1.607	-1.785	-1.859	1.318

[a] ADF(r) denotes the Augmented Dickey-Fuller test with r lags. The 1%, 5%, and 10% critical values are -4.060, -3.459, and -3.155, respectively. These values are calculated from Table 1 of Mackinnon (1991) for T=92.

[b] The 1%, 5%, and 10% critical values for J(1,5) are 0.123, 0.295, and 0.452, respectively. When the J(1,5) statistic is smaller than these values, the null hypothesis of difference stationarity is rejected.

Table 3
Cointegrating Regression Results

ϵ	d	H(0,1)	H(1,2)	H(1,3)	H(1,4)
2.537		1.793	1.762	8.941	9.026
(0.257)		[0.181]	[0.184]	[0.011]	[0.029]
2.208	-0.259	0.741	0.450	1.216	1.689
(0.100)	(0.042)	[0.389]	[0.502]	[0.545]	[0.639]

[a] The second column presents the estimated coefficient of the intercept dummy variable for the period of the bubble economy.

[b] The numbers in parentheses are standard errors.

[c] H(0,1) denotes a χ^2 test statistic with one degree of freedom for the deterministic cointegration restriction. H(1,q) denotes a χ^2 test statistic with $q-1$ degrees of freedom for stochastic cointegration. The numbers in square brackets are asymptotic p-values.

Table 4
GMM Estimates of the Intertemporal Elasticity of Substitution
(The Two-Good Model)

	ϵ	a	β	σ	J_T
(1)	2.208	7.595	0.999	0.771	1.375
				(0.127)	[0.711]
(2)	2.208	7.595	0.998	0.900	1.610
				(0.173)	[0.657]
(3)	2.208	7.595	0.997	1.072	1.911
				(0.246)	[0.591]
(4)	2.208	7.920	0.999	0.762	1.355
				(0.126)	[0.716]
(5)	2.008	7.595	0.999	0.746	1.325
				(0.121)	[0.723]

[a] The numbers in parentheses are standard errors.

[b] J_T denotes Hansen's J test with three degrees of freedom. The numbers in square brackets are asymptotic p-values.

Table 5
GMM Estimates of the Intertemporal Elasticity of Substitution
(The One-Good Model)

	β	σ	J_T
(A) Nondurables ($\sigma = \epsilon$)			
(1)	0.999	0.411	0.670
		(0.084)	[0.880]
(2)	0.998	0.511	0.824
		(0.130)	[0.844]
(3)	0.997	0.667	1.101
		(0.220)	[0.777]
(B) Nondurables+Service Flow			
(1)	0.999	1.202	8.770
		(0.068)	[0.033]
(2)	0.998	1.311	9.628
		(0.073)	[0.022]
(3)	0.997	1.469	11.567
		(0.087)	[0.009]
(C) Total Consumption			
(1)	0.999	1.200	12.353
		(0.203)	[0.006]
(2)	0.998	1.539	12.257
		(0.296)	[0.007]
(3)	0.997	2.055	11.659
		(0.461)	[0.009]

[a] The numbers in parentheses are standard errors.

[b] J_T denotes Hansen's J test with three degrees of freedom. The numbers in square brackets are asymptotic p-values.

Appendix

In this appendix, we examine whether the growth rate of estimated marginal utility, mu_{t+1}/mu_t, is stationary. To test the null hypothesis of level stationarity, we use both Park's (1990) G(p,q) test and Kwiatkowski, Phillips, Schmidt, and Shin's (1992) test (KPSS test). The estimated marginal utility is calculated using the baseline results in Table 4. Table A reports the test results. As this table shows, the null hypothesis of stationarity cannot be rejected for the growth rate of the estimated marginal utility.

Table A

Tests for Stationarity of the Growth Rate of Estimated Marginal Utility

KPSS(1)	KPSS(4)	KPSS(7)	G(0,1)	G(0,2)	G(0,3)
(1) $\beta = 0.999 : \sigma = 0.771, \epsilon = 2.208$					
0.054	0.085	0.095	0.321	1.183	3.157
			[0.571]	[0.553]	[0.368]
(2) $\beta = 0.998 : \sigma = 0.900, \epsilon = 2.208$					
0.055	0.089	0.101	0.411	1.182	3.025
			[0.521]	[0.554]	[0.388]
(3) $\beta = 0.997 : \sigma = 1.072, \epsilon = 2.208$					
0.059	0.097	0.111	0.549	1.207	2.881
			[0.459]	[0.547]	[0.410]

[a] KPSS(p) denotes the KPSS test with p lags. The 1 %, 5 %, and 10 % critical values are 0.739, 0.463, and 0.347, respectively. These are taken from Kwiatkowski *et al.* (1992).

[b] G(0,q) denotes a χ^2 test statistic with q degrees of freedom. The numbers in square brackets are asymptotic p-values.