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Measurement of Intertemporal Substitution under Nonseparable and Nonhomothetic Utility: The Case of Import Demand *

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Abstract

When estimating the intertemporal elasticity of substitution (IES), although adopting a constant elasticity of substitution utility function may lead to a more plausible estimate, the assumption of homotheticity obviously contradicts the data. This paper introduces a model with both nonseparability and nonhomotheticity of preferences and extends the use of the existing two-step approach that combines cointegration with generalized methods of moments. Applying our approach to U.S. import demand data, we demonstrate that the IES estimates are significant and similar to those found by more recent studies. We argue, however, that this evidence cannot be obtained from the existing approach.

Key words: Intertemporal elasticity of substitution, Nonhomothetic preferences, Habit formation, Cointegration, Generalized Method of Moments

JEL classification numbers: C22, E21, F41

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1 Introduction

In open-economy analysis, the real exchange rate, the real interest rate and the terms of trade affect the optimal allocation of consumption, and thereby influence the adjustment of the economy to economic shocks, with consequent policy implications. On this basis, most modern international macroeconomic models have allowed for both intertemporal and *intra*temporal relative price changes, typically by assuming a constant elasticity of substitution and constant relative risk aversion (CES–CRRA) utility function (see, e.g., Obstfeld and Rogoff (1996)).

However, in empirically analyzing the intertemporal aspect of import demand, many studies have assumed additively separable preferences in domestic and imported nondurable goods (see, e.g., Ceglowski (1991), Clarida (1994, 1996), Amano and Wirjanto (1996), de la Croix and Urbain (1998) and Nishiyama (2002)). As Ostry and Reinhart (1992) and Ogaki and Reinhart (1998a,b) have emphasized, ignoring intratemporal relative price changes by assuming separability can seriously bias the estimates of preference parameters. Nevertheless, there are few empirical studies based on models with nonseparability between goods in this field.

An exception is Amano et al. (1998), who attempted to fill the gap by developing a model with CES–CRRA preferences. However, it is known that such an approach to allow for nonseparability generally creates two problems, which explains the lack of empirical investigation. First, we need to use a two-step procedure that combines cointegration with generalized methods of moments (GMM) estimation of the Euler equation to estimate the IES. However, when the stationary error term introduced into the first-step cointegrating regression is interpreted as arising from preference shocks and measurement errors, as has been

done in some of the studies cited above, the second-step GMM estimators are inconsistent because such methods affect the specification of the Euler equation.[1] Hence, in utilizing a two-step procedure, it is particularly important to clarify the rationale for adding the stationary error term to the regression,[2] although Amano et al. (1998) did not focus on this econometric issue.

Second, we must address the theoretical issue of CES-type utility functions being homothetic. This assumption, in the context of import demand, implies that both domestic and imported goods are neither necessities nor luxuries. According to some empirical studies,[3] imported goods are likely to be luxury goods. For this reason, it is often pointed out that the adoption of CES–CRRA preferences may bias the estimation of the IES,[4] even if it complements the main shortcoming of previous studies. However, no study has attempted to estimate the IES by using a model that controls for the effect of the homotheticity assumption, at least in the present context.

The main purpose of this paper is to provide an empirically feasible solution to these econometric and theoretical issues. Specifically, we consider using the addilog utility function with time-nonseparability of consumption used in Ogaki and Park (1998). We raise this function to the power of 1 minus the reciprocal of the IES to allow for both nonseparability and

[1] A similar problem arises when conducting the specification test that compares cointegration regression estimates with GMM estimates. See, e.g., Ogaki and Park (1998) and Nishiyama (2002) for detailed discussions.

[2] For CES–CRRA preferences, Ogaki and Reinhart (1998a,b) and Okubo (2003a) developed two-step procedures on this basis. However, their approach has not been applied to import-demand analysis. The reason for this is given in the next paragraph.

[3] See, e.g., Ait-Sahalia et al. (2004) for a recent discussion.

[4] As Vissing-Jørgensen (2002) showed, since the use of aggregate data might itself bias estimation of the IES, it is appropriate to use a micro-data. However, as Browning and Crossley (2000) have pointed out, most micro-data studies use single-good models with CRRA preferences because of their tractability and because of data availability. If more suitable panel data are available, the type of preferences that should be assumed, to some extent, will rely on empirical findings based on aggregate data. Given the lack of empirical investigation, it is still meaningful to evaluate the relative importance of factors such as nonseparability and nonhomotheticity by using aggregate data, at least in the early stages of analysis.

nonhomotheticity of preferences. The principal advantage of using this utility function is that since it incorporates as special cases most of the utility functions used in existing empirical studies, it enables a formal test of homotheticity. On the basis of this extended model, we derive a cointegration restriction that implies a cointegration relationship between domestic goods, imported goods, and their relative prices, and then extend the use of the two-step procedure. This paper represents the first to attempt to evaluate intertemporal substitution in import demand by incorporating both nonseparability and nonhomotheticity.[5]

In the first step of our procedure, we estimate the curvature parameters governing domestic and imported consumption and test the null hypothesis of homothetic preferences. This step can be interpreted as a simple generalization of the first-step estimation procedure used in Ogaki and Reinhart (1998a,b). In the second step, GMM is applied to the Euler equation incorporating time-and goods-nonseparability and nonhomotheticity. This equation represents an extension of the Euler equations estimated by Ferson and Constantinides (1991), Cooley and Ogaki (1996) and Ogaki and Park (1998). In this second step, we estimate the IES for a composite of domestic and imported goods. Consequently, our estimator of the IES is robust to various factors that affect intratemporal and intertemporal consumption choice.

In our empirical work, we apply our two-step procedure to data on U.S. import demand. In the first step, we show that the assumption of homotheticity is rejected by the formal test. In the second step, using GMM estimation, we obtain IES estimates that are significantly different from zero. Our estimates are similar to those of Ogaki and Reinhart (1998a). However, we argue that these estimates cannot be obtained from models that assume homotheticity

[5] Under the assumption of time-and goods-separability, Atkeson and Ogaki (1996) considered allowing for nonhomotheticity by incorporating fixed subsistence levels and proposed a similar estimation method based on both aggregate and panel data.

together with nonseparability.

The rest of this paper is organized as follows. In section 2, we explain our theoretical framework. In section 3, we describe our two-step estimation procedure. In section 4, we examine the time-series properties of the data used for analysis. In section 5, we present our estimation and test results. Section 6 contains concluding remarks.

2 The Model

Suppose that a representative consumer maximizes expected lifetime utility

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

subject to the following budget constraint:

$$A_{t+1} = (1 + r_t) A_t + Y_t - \sum_{i=d,m} P_{it} C_{it}, \tag{2}$$

where $E_t[\cdot]$ is the expectations operator conditional on the information available at time t, β is the subjective discount factor, σ is the IES, C_{dt} is the consumption of domestic nondurable goods in period t, C_{mt} is the consumption of imported nondurable goods in period t, A_t is the asset holding of the representative consumer, r_t is the one-period interest rate on assets between $t-1$ and t, Y_t is the labor income in period t, and P_{it} is the price of goods i $(i = d, m)$ in period t.

We assume that the period utility function takes the following form:

$$u(t) = \left[k \frac{C_{dt}^{*1-\alpha}}{1-\alpha} + \frac{C_{mt}^{*1-\gamma}}{1-\gamma} \right]^{1-\frac{1}{\sigma}}, \tag{3}$$

where $\alpha > 0$ and $\gamma > 0$ are the curvature parameters,[6] $k > 0$ is a scaling factor, and C_{it}^* is

[6] When there is habit formation, it should be noted that the inverses of these parameters cannot be interpreted as the IES.

defined as

$$C_{it}^{*} = C_{it} + \theta_i C_{it-1}, \quad i = d, m, \tag{4}$$

so that, depending on the values of θ_i, this model incorporates habit formation. This utility function is the addilog utility function of Ogaki and Park (1998) raised to the power 1 minus the reciprocal of the IES. As is easily verified from (3), the marginal utilities of domestic and imported goods consumption are not independent of each other, and homotheticity is not imposed a priori. This slight generalization of the utility function incorporates as special cases most of the utility functions examined by previous empirical studies.

If $1/\sigma = 0$ (i.e., if separability in goods is assumed), this utility function is equivalent to those studied by Amano and Wirjanto (1996) and de la Croix and Urbain (1998). In addition, if $\theta_i = 0$ (i.e., if no habit formation is assumed), the utility function is equivalent to those estimated by Ceglowski (1991), Clarida (1994), Amano and Wirjanto (1996) and Nishiyama (2002).[7] Alternatively, under the assumption that $1/\sigma \neq 0$, and supposing that $\theta_i = 0$ and $\alpha = \gamma$, we obtain the CES-type utility function (i.e., a nonseparable but homothetic utility function), used by Ostry and Reinhart (1992), Ogaki et al. (1996), Amano et al. (1998) and Okubo (2003a). Moreover, if $\theta_d = 0$ and $\theta_m \neq 0$ (or $\theta_d \neq 0$ and $\theta_m = 0$), this functional form incorporates one similar to the CES-type utility function introduced by Ogaki and Reinhart (1998a,b).

Solving the utility-maximization problem of the representative consumer, the Euler equa-

[7] More precisely, for this specification of utility function, Clarida (1994) and Amano and Wirjanto (1996) further assumed unknown preference shocks, which the authors interpreted as stationary error terms. As we mentioned in the introduction, in the two-step approach, this assumption causes the second-step estimator to be inconsistent. Therefore, we do not make such an assumption in this paper. However, Nishiyama (2002) proposed the cross-Euler equation approach, which introduces a stationary error term without affecting the Euler equation. Since our paper assumes time-nonseparability, we need not use this approach.

tions that characterize the intertemporal allocation of consumption are

$$E_t\left[\beta(1+r_{t+1})\frac{P_{it}}{P_{it+1}}\frac{\mu_i(t+1)+\beta\theta_i\mu_i(t+2)}{\mu_i(t)+\beta\theta_i\mu_i(t+1)}\right]=1,\quad i=d,m, \tag{5}$$

and the intraperiod first-order condition is given by

$$\begin{aligned}\frac{P_{dt}}{P_{mt}}=\frac{\partial U/\partial C_{dt}}{\partial U/\partial C_{mt}} &= \frac{E_t[\partial u(t)/\partial C_{dt}+\beta\partial u(t+1)/\partial C_{dt}]}{E_t[\partial u(t)/\partial C_{mt}+\beta\partial u(t+1)/\partial C_{mt}]}\\ &= \frac{kE_t[\mu_d(t)+\beta\theta_d\mu_d(t+1)]}{E_t[\mu_m(t)+\beta\theta_m\mu_m(t+1)]}\\ &= \frac{kE_t[C_{dt}^{*-\alpha}+\beta\theta_d C_{dt}^{*-\alpha}\{\mu_d(t+1)/\mu_d(t)\}]}{E_t[C_{mt}^{*-\gamma}+\beta\theta_m C_{mt}^{*-\gamma}\{\mu_m(t+1)/\mu_m(t)\}]},\end{aligned} \tag{6}$$

where

$$\mu_d(t)=C_{dt}^{*-\alpha}\left[k\frac{C_{dt}^{*1-\alpha}}{1-\alpha}+\frac{C_{mt}^{*1-\gamma}}{1-\gamma}\right]^{-\frac{1}{\sigma}}, \tag{7}$$

$$\mu_m(t)=C_{mt}^{*-\gamma}\left[k\frac{C_{dt}^{*1-\alpha}}{1-\alpha}+\frac{C_{mt}^{*1-\gamma}}{1-\gamma}\right]^{-\frac{1}{\sigma}}. \tag{8}$$

The last equality of (6) is obtained by dividing the numerator and denominator of the second equality by $\mu_d(t)$ and rearranging terms. Multiplying both sides of equation (6) by $C_{mt}^{-\gamma}/C_{dt}^{-\alpha}$ yields

$$\frac{P_{dt}}{P_{mt}}\frac{C_{mt}^{-\gamma}}{C_{dt}^{-\alpha}}=\frac{kE_t[(C_{dt}^*/C_{dt})^{-\alpha}+\beta\theta_d(C_{dt}^*/C_{dt})^{-\alpha}\{\mu_d(t+1)/\mu_d(t)\}]}{E_t[(C_{mt}^*/C_{mt})^{-\gamma}+\beta\theta_m(C_{mt}^*/C_{mt})^{-\gamma}\{\mu_m(t+1)/\mu_m(t)\}]}, \tag{9}$$

where the growth rates of the marginal utilities are

$$\frac{\mu_d(t+1)}{\mu_d(t)}=\left[\frac{C_{dt+1}^*}{C_{dt}^*}\right]^{-\alpha}\left[\frac{k(1-\gamma)C_{dt+1}^{*1-\alpha}+(1-\alpha)C_{mt+1}^{*1-\gamma}}{k(1-\gamma)C_{dt}^{*1-\alpha}+(1-\alpha)C_{mt}^{*1-\gamma}}\right]^{-\frac{1}{\sigma}}, \tag{10}$$

$$\frac{\mu_m(t+1)}{\mu_m(t)}=\left[\frac{C_{mt+1}^*}{C_{mt}^*}\right]^{-\gamma}\left[\frac{k(1-\gamma)C_{dt+1}^{*1-\alpha}+(1-\alpha)C_{mt+1}^{*1-\gamma}}{k(1-\gamma)C_{dt}^{*1-\alpha}+(1-\alpha)C_{mt}^{*1-\gamma}}\right]^{-\frac{1}{\sigma}}. \tag{11}$$

That is, the right-hand side of (9) is a function of C_{it}^*/C_{it} and $\mu_i(t+1)/\mu_i(t)$, and $\mu_i(t+1)/\mu_i(t)$ can be expressed as a function of C_{it+1}^*/C_{it}^* and the growth rate of the composite good $[kC_{dt}^{*1-\alpha}/(1-\alpha)+C_{mt}^{*1-\gamma}/(1-\gamma)]$.

Since empirical analysis is our main concern, we make the additional assumptions that $\ln(C_{dt})$, $\ln(C_{mt})$ and $\ln(P_{dt}/P_{mt})$ are difference stationary, following previous works on import demand. Given this assumption, it can be shown that C^*_{it}/C_{it} and C^*_{it+1}/C^*_{it} are stationary for $i = d, m$ (see Cooley and Ogaki (1996) and Ogaki and Park (1998)), which implies that the first terms in the numerator and denominator on the right-hand side of (9) are stationary. However, it is possible that the growth rate of marginal utility $\mu_i(t+1)/\mu_i(t)$ is nonstationary, because the assumption of difference stationarity is not sufficient to ensure that the growth rate of the composite good is stationary. As Ogaki and Reinhart (1998a) and Okubo (2003a) explain, this nonstationarity problem arises whenever nonseparability of preferences is assumed. However, Ogaki and Reinhart (1998a) and Okubo (2003b) empirically demonstrated that this is unlikely to cause a serious problem, at least in the context of using U.S. and Japanese data. Hence, as did these authors, we also assume that the growth rate of the composite good is stationary, and subsequently test the validity of this assumption using the estimated values of the composite good. Under these assumptions, $P_{dt}C_{mt}^{-\gamma}/P_{mt}C_{dt}^{-\alpha}$ is stationary because the right-hand side of (9) is stationary, which, after taking logarithms of both sides of (9), yields the cointegration restriction that $(\ln(P_{dt}/P_{mt}), \ln C_{dt}, \ln C_{mt})'$ is cointegrated with the cointegrating vector $(1, \alpha, -\gamma)'$.

3 Estimation Procedures

We have shown that the intraperiod first-order condition implies a cointegration restriction. This restriction can be used to estimate and test the model. In this paper, we essentially follow the method developed by Ogaki and Reinhart (1998a,b). However, our approach involves a few modifications. In this section, we describe an extended version of the existing

two-step procedure that combines the cointegration approach with GMM estimation.

3.1 Cointegration and Tests for Homotheticity of Preferences

From equation (9), we obtain the following cointegrating regression:

$$\ln(P_{dt}/P_{mt}) = \ln(k) - \alpha \ln(C_{dt}) + \gamma \ln(C_{mt}) + u_t, \tag{12}$$

where u_t is a stationary error term. This relationship enables us to formulate model specification tests as well as an approach to estimating the preference parameters (α and γ) and the scaling factor (k). As a general case, suppose that the consumption and relative price series are difference stationary with drift. Since our model does not generate a trend term in the cointegrating regression, it implies that the cointegrating vector eliminates both the deterministic trends arising from drift and the stochastic trends. Therefore, if our specification of nonseparable and nonhomothetic preferences is correct, the null hypothesis of deterministic cointegration (i.e., stochastic cointegration with the deterministic cointegration restriction) will not be rejected, and the null hypothesis of homotheticity (i.e., $\alpha = \gamma$) will be rejected.

To implement these specification tests in practice, we first apply Park's (1992) canonical cointegrating regression (CCR) to equation (12) and then use Park's (1990) H(p,q) tests. H(p,q) test statistics are Wald tests for superfluous deterministic trends in the test equation. Depending on the types of deterministic trends included in the equation, the orders of p and q are determined. Following the standard cointegration approach, H(0,1) test is used to test the deterministic cointegration restriction, and H(1,q) tests are used to test stochastic cointegration. To test the null hypothesis of homotheticity, we employ the K statistic, which is a Wald test of the hypothesis $\alpha = \gamma$ in the cointegrating regression. Under this hypothesis, the K statistic is asymptotically χ^2 distributed with one degree of freedom because of the

asymptotic normality of the CCR estimators. In the first step of our procedure, we evaluate the validity of the model based on these three kinds of test statistic.

3.2 Estimation of the Intertemporal Elasticity of Substitution

In the second step, GMM estimation is based on the Euler equation (5), which implies $E_t[e^0_{dt}] = 0$,[8] where

$$e^0_{dt} = \beta(1+r_{t+1})\frac{P_{dt}}{P_{dt+1}}\{\mu_d(t+1) + \beta\theta_d\mu_d(t+2)\} - \{\mu_d(t) + \beta\theta_d\mu_d(t+1)\}. \tag{13}$$

However, this disturbance term cannot be used as the basis for GMM estimation because it involves nonstationary variables such as C^*_{dt} and C^*_{mt}. Hence, the regularity conditions for GMM estimators are violated. Following Ferson and Constantinides (1991), if we are assuming that preferences are separable between the two goods, this problem can be avoided by defining $e_{dt} = e^0_{dt}/C^{*-\alpha}_{dt}$. However, under the assumption of nonseparable preferences, this scaled disturbance still includes nonstationary terms such as $[kC^{*1-\alpha}_{dt}/(1-\alpha)+C^{*1-\gamma}_{mt}/(1-\gamma)]$.

To achieve the stationarity required for GMM, we transform the disturbance so that it comprises growth rates of marginal utilities. We achieve this by dividing the disturbance e^0_{dt} by $\mu_d(t)$ as follows:

$$\begin{aligned} e^*_{dt} &= e^0_{dt}/\mu_d(t) \\ &= \beta(1+r_{t+1})\frac{P_{dt}}{P_{dt+1}}\left\{\frac{\mu_d(t+1)}{\mu_d(t)} + \beta\theta_d\frac{\mu_d(t+2)}{\mu_d(t)}\right\} - \left\{1 + \beta\theta_d\frac{\mu_d(t+1)}{\mu_d(t)}\right\}. \end{aligned} \tag{14}$$

Since $\mu_d(t)$ is a function of C^*_{dt} and C^*_{mt}, it is in the information set available at time t, so that $E_t[e^*_{dt}] = 0$. When the real return $(1+r_{t+1})P_{dt}/P_{dt+1}$ is assumed to be stationary, as in the

[8] Given that the intraperiod first-order condition (6) holds, for estimation, we can concentrate on the Euler equation for one of the two goods. Following convention, as have Ogaki et al. (1996), Amano et al. (1998) and Ogaki and Reinhart (1998a,b), we focus on the Euler equation for domestic nondurable consumption, which facilitates comparison with previous studies.

standard GMM approach, and when the stationarity of marginal utility growth is assumed, the disturbance e^*_{dt} is a function of stationary variables. Given $1/\sigma = 0$, equation (14) can be reduced to the GMM disturbance used by Cooley and Ogaki (1996) and de la Croix and Urbain (1998). In this sense, the disturbance term in our model is a generalization of those in previous studies.

Let $\hat{e}^*_{dt}$ denote the disturbance term into which the estimates of α, γ and k from the cointegrating regression in the first step were substituted. Let z_t be a vector of instrumental variables included in the information at time t. In the second step of our procedure, we apply GMM to the disturbance term $\hat{u}^*_t = z_t \hat{e}^*_t$, and thereby estimate the remaining parameters. In this model, since the disturbance term $\hat{u}^*_t$ involves both C_{it+1} and $C_{it+2}\,(i = d, m)$, it is in the information at time $t + 2$. Therefore, as in the estimation of the standard model with time-nonseparability, we allow $\hat{u}^*_t$ to have a moving average of order one (MA(1)) structure. This two-step procedure does not affect the asymptotic properties of the GMM estimators and test statistics due to super-consistency of the cointegrating regression estimators.

4 Data

We use seasonally adjusted U.S. quarterly data on domestic and imported nondurables expenditures. The corresponding implicit deflators are used for the prices of the consumption series, and the three-month Treasury bill rate is used for nominal interest rates. The sample period covers 1967:1 to 1994:3 (111 observations). Our measure of nondurable consumption is food plus non-oil nondurables. The required rate of return is defined as the real return plus a constant risk premium of 2% per quarter. The data are the same as those used in de la Croix and Urbain (1998) and are constructed by following the definitions of Ceglowski

(1991). Since the details of data construction and data sources are in their papers, this section summarizes the statistics that serve our purposes and reports additional unit root test results.

Table 1 presents selected summary statistics of the data. The first two columns of the table report the shares of domestic and imported goods in total expenditure (defined as the sum of the two expenditures). The next two columns report the log of the price of each good for the sample period. This table shows there have been changes in the structure of consumption expenditures over the sample period. First, the budget share of imported goods has increased by 53.7% (i.e., 38.9%→59.8%), while the domestic goods share has fallen by 34.2% (i.e., 61.1%→40.2%). Second, the (log) prices of domestic and imported goods have increased by 1.34 and 1.45, respectively.[9]

The budget share of imported goods reported in this paper is large compared to those reported by other studies. For example, Amano et al. (1998) report budget shares of around 1% in 1967 and around 4.7% in 1993 for imported goods. This difference between the figures is probably due to different data definitions and different measures of total consumption.[10] However, more importantly, the above observable changes in the budget shares contradict the implication of homothetic preferences that all expenditure elasticities are unity. As explained in the introduction, given this finding, the assumption of homotheticity may be a source of bias in estimation of the IES.

On the basis of this data set, de la Croix and Urbain (1998) and Nishiyama (2002) have already found that the null hypothesis of difference stationarity cannot be rejected for

[9] Real total consumption expenditure rose by 152.5% over the sample period.

[10] This paper does not attempt to unify the findings of these previous studies. Thus, we proceed the following discussions as one possible case derived from de la Croix and Urbain's data set.

$\ln(C_{dt})$, $\ln(C_{mt})$, and $\ln(P_{mt}/P_{dt})$. To confirm these findings, we conduct alternative tests of the null hypothesis of difference stationarity. Before conducting formal tests, it is useful to examine the changes in the three series over time. Plots of these series for our sample period are shown in Figure 1. This figure shows that while $\ln(C_{dt})$ and $\ln(C_{mt})$ have clear upward trends, $\ln(P_{dt}/P_{mt})$ appears to exhibit a slightly downward trend, suggesting that these series may be difference stationary processes with drifts.

Table 2 reports the test results for the null hypothesis of difference stationarity. We use three recent test statistics proposed by Ng and Perron (2001). These are a modified version of the Phillips and Perron (1988) Z_α test, a modified version of the Elliott et al. (1996) DF-GLS test and a modified feasible point optimal test. The tests are referred to as the $\bar{M}Z_\alpha^{GLS}$ test, the DF^{GLS} test, and the $\bar{M}P_T^{GLS}$ test, respectively. We use the modified Akaike information criterion. The table shows that these tests fail to reject the null hypothesis of difference stationarity for all variables, even at the 10% significance level. Thus, the assumptions required for deriving the cointegration restriction are not rejected.

5 Empirical Results

We now examine the empirical validity of our model specification. Table 3 reports the CCR estimates for the cointegrating regression and the test statistics described in section 3. Panel A of the table indicates that the constant term and the curvature parameters α and γ are statistically significant and have theoretically expected signs. The implied estimate for the scaling factor k is 1.080. The implied values of $1/\alpha$ and $1/\gamma$ are 0.270 and 0.754, respectively. Since the H(0,1) test statistic is not significant at conventional significance levels, the deterministic cointegration restriction is not rejected. The H(1,2) test statistic is significant at the

10% level, but is not significant at the 5% level. The H(1,3) and H(1,4) test statistics are not significant even at the 10% level. Hence, overall the H(1,q) tests do not reject the null hypothesis of stochastic cointegration. In addition, the hypothesis that $\alpha = \gamma$, which implies that preferences are homothetic, is strongly rejected by the K test. This is consistent with our observations concerning the budget share changes discussed in section 4.

To confirm the effect of the homotheticity assumption on the estimation and test results, we estimated the cointegrating regression with the restriction $\alpha = \gamma$. Panel B of the table shows that the CCR estimate of the curvature parameter is not significant, and the H(1,2), H(1,3), and H(1,4) statistics do not support stochastic cointegration. Thus, the specification that imposes homotheticity is strongly rejected. In summary, the results show that the cointegration restriction derived from our model incorporating both nonseparability and nonhomotheticity is supported by the data.

Table 4 reports the estimation results of the Euler equation based on the second step of GMM. We used several sets of instrumental variables with various lags. The instrument sets, which are listed in the note to Table 4, were chosen by following the literature on the GMM estimation of models with nonseparability in goods and models with habit formation. To allow for the MA(1) structure of the GMM disturbance term, we calculated the weighting matrix under the assumption that the disturbance follows an MA(1) process with an unknown coefficient. Each GMM estimation was iterated five times.

When θ_d and θ_m were estimated together with the parameters β and σ, we encountered convergence problems, which prevented us from obtaining these estimates. Therefore, following Ogaki and Reinhart (1998a), we report results based on fixed values of θ_d and θ_m. According to de la Croix and Urbain (1998), when using the present data set, habit formation appears

significant and more important for domestic goods. Hence, we adopt three pairs of estimates for θ_d and θ_m, which are available from their paper: $(\theta_d, \theta_m) = (-0.74, -0.42), (-0.55, -0.23)$ and $(-0.53, -0.36)$. For our purposes, this approach is convenient for evaluating the robustness of the IES estimates in two respects, the choice of instrument sets for the same degree of habit formation, and the difference in habit formation for the same instrument sets. Panels A, B, and C of Table 4 report results based on these three pairs.

The following is apparent from the panels of Table 4. First, on the basis of Hansen's J test of the overidentifying restrictions, the evidence against our model is not overwhelming. In Panel A, there are eight rejections from 15 tests at the 5% significance level, whereas in Panels B and C, there are four and three rejections, respectively. Thus, two-thirds of the 45 tests do not reject the model at conventional significance levels. Second, the discount factor β is significantly estimated, at between 0.95 to 0.98. Third, and more importantly, the point estimates of the IES are significantly different from zero in almost all cases. In Panel A, the estimates of σ range from 0.5 to 0.6. In Panel B, most estimates range from 0.3 to 0.7. Similarly, those in Panel C range from 0.3 to 0.8. These results suggest that the estimates of σ are in a relatively narrow range, and are not very sensitive to the choice of (θ_d, θ_m).[11] Further, our IES estimates are similar to those from recent studies, particularly Ogaki and Reinhart (1998a), whose estimates of IES were between 0.3 and 0.5, using U.S. quarterly data based on the CES–CRRA utility function described in section 2. In addition, as is clear from the small standard errors for σ, the separability assumption $(1/\sigma = 0)$ is rejected, except in a few cases.

[11] However, this does not mean that the introduction of habit formation is not important. Rather, this implies that nonseparability in goods is useful for controlling for the effect on the IES estimates of changes in the relative importance of habits between domestic and imported goods.

Comments on the second-step GMM results are in order because our results may appear to contrast with those from the standard GMM approach. As Stock and Wright (2000), among others, have pointed out, the model diagnosis based on Hansen's J-test statistic may need to be interpreted with caution. For example, it is often the case that although Hansen's J test does not reject a model, the estimated preference parameters are statistically significant but have theoretically implausible signs.[12] However, this does not apply to our results. On the other hand, unlike the simple GMM approach, our second-step GMM estimation uses information on stochastic and deterministic trends by incorporating the first-step estimates. One would expect this to reduce estimation problems. That is, use of the two-step procedure avoids difficulties apparent in the literature. Given that our specification passed the J test and the specification tests from the first-stage, our results seem unaffected by the problems that have beset the standard GMM approach.

As we noted in section 2, stationarity of the growth rate of the composite good is important in the context of the two-step procedure. Given different values of (θ_d, θ_m), we obtain three series for the growth rate. These are plotted in Figure 2. Clearly, these cannot be regarded as trend stationary processes. We confirm this by using Park's (1990) G(0,q) test.[13] Table 5 reports test results for the null hypothesis of level stationarity for the growth rate of the estimated composite good. Our test results indicate that the null hypothesis of stationarity cannot be rejected for all cases. Thus, as in Ogaki and Reinhart (1998a) and Okubo (2003b), we find that the nonstationarity problem does not arise empirically.

[12] Nishiyama (2002) encountered this problem when estimating an Euler equation for domestic nondurables using the standard GMM approach.

[13] Following Kahn and Ogaki's (1992) Monte Carlo simulations, we use Andrews's (1991) quadratic spectral kernel with the automatic bandwidth parameter estimator based on AR(1).

6 Concluding Remarks

In empirical analysis of a model with nonseparable preferences between goods, two problems have been identified by the existing literature. First, in the estimation of preference parameters, the derivation of the stationary error term in the cointegrating regression in the first step should not affect estimation in the second step. Second, even if this can be achieved, and one allows for the *intra*temporal substitution effect, because homothetic preferences are assumed, estimates of the intertemporal elasticity of substitution (IES) may be biased. This trade-off is a typical problem that arises when investigating intertemporal substitution in import demand.

In this paper, we introduced a two-good model with nonseparable and nonhomothetic preferences in domestic and imported nondurable goods, and proposed an alternative two-step procedure for estimating preference parameters. We applied this extended approach to U.S. import demand data, and attempted to estimate the IES. We obtained statistically significant IES estimates of around 0.3 to 0.8. We also found empirical evidence against the separability and homotheticity of preferences between domestic and imported goods.

Our IES estimates, obtained from a model with nonseparability and nonhomotheticity, are similar to those reported by recent studies that have found a statistically significant IES. However, this does not mean that the existing approach is applicable. In our empirical work, since the cointegrating regression imposing the restriction $\alpha = \gamma$ was strongly rejected, the first-step estimates, which are required for the second-step GMM estimation, could not be obtained. Thus, our estimation results still provide questions about the existing IES estimates obtained on the basis of assuming homotheticity. In simulation studies, can such IES estimates be adopted as an approximation? Conversely, solely on the basis of the homotheticity

issue, are such IES estimates unreliable?

At least in the context of empirical analysis of an open economy, the relative importance of the factors that affect intratemporal and intertemporal consumption choices depend on the economic environment and that economy's international position. Thus, further investigation of our model using data from other countries is of interest for future research.

In this paper, we assumed that preferences are separable between nondurable and durable goods. As Ogaki and Reinhart (1998a,b) showed, when U.S. data are used, this assumption affects the estimates of the IES. However, their model also imposed homotheticity over nondurable and durable goods. Thus, to further investigate the effect of the homotheticity assumption on the IES estimates, it would be interesting to estimate a modified version of our model using their data. This extension is currently being undertaken by the author.

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Table 1
Budget Share and Price Changes over the Sample Period

	Budget Share (%)		Log Price	
Good	1967:1	1994:3	1967:1	1994:3
Domestic nondurables	61.1	40.2	-1.10	0.24
Imported nondurables	38.9	59.8	-1.25	0.20

Note: The budget share is defined as the ratio of each good to the sum of domestic and imported goods.

Table 2
Unit Root Test Results

Variable	k_{maic}	$\bar{M}Z_{\alpha}^{GLS}$	DF^{GLS}	$\bar{M}P_{T}^{GLS}$
$\ln(C_{dt})$	0	-7.788	-1.945	12.012
$\ln(C_{mt})$	1	-10.803	-2.430	8.451
$\ln(P_{dt}/P_{mt})$	1	-3.787	-1.381	23.510

Notes: k_{maic} denotes the number of lags determined by the modified Akaike information criterion, where an upper bound on the lag length is set at the integer part of $12(T/100)^{1/4}$. The test regressions include both a constant and a trend term. The factor $\bar{c}$ required to construct the GLS detrended series is set at -13.5, following Elliott et al. (1996). Critical values at the 1%, 5%, and 10% significance levels are -23.8, -17.3, and -14.2 for the $\bar{M}Z_{\alpha}^{GLS}$ test, -3.42, -2.91, and -2.62 for the DF^{GLS} test, and 4.03, 5.48, and 6.67 for the $\bar{M}P_{T}^{GLS}$ test. These critical values are from Table I of Ng and Perron (2001).

Table 3
Estimation Results from the Cointegrating Regression

$\ln(k)$	α	γ	H(0,1)	H(1,2)	H(1,3)	H(1,4)	K
Panel A: $\ln(P_{dt}/P_{mt}) = \ln(k) - \alpha \ln(C_{dt}) + \gamma \ln(C_{mt}) + u_t$							
0.077	3.702	1.326	0.042	3.419	3.897	4.752	215.654
(0.011)	(0.251)	(0.093)	[0.838]	[0.064]	[0.142]	[0.191]	[0.000]
Panel B: $\ln(P_{dt}/P_{mt}) = \ln(k) + \alpha \ln(C_{mt}/C_{dt}) + u_t$							
0.010	0.033		2.260	14.019	15.462	23.273	
(0.055)	(0.137)		[0.133]	[0.000]	[0.000]	[0.000]	

Notes: Park's (1992) canonical cointegrating regression estimates are based on the quadratic spectral kernel and the VAR(1) prewhitening technique of Andrews and Monahan (1992). Standard errors are in parentheses. H(0,1) is a χ^2 test statistic for the null hypothesis of the deterministic cointegration restriction. H(1,2), H(1,3), and H(1,4) are χ^2 test statistics for the null hypothesis of stochastic cointegration. K is a χ^2 test statistic for the null hypothesis of homotheticity $\alpha = \gamma$. P-values are in square brackets.

Table 4
Generalized Method of Moments Results

Instruments	Lag	β	s.e.	σ	s.e.	J_T	p-value	df
Panel A: $\theta_d = -0.74$, $\theta_m = -0.42$								
Inst1	(-1)	0.978	(0.006)	0.606	(0.028)	3.582	[0.058]	1
	(-2)	0.973	(0.004)	0.583	(0.014)	6.637	[0.084]	3
	(-3)	0.975	(0.003)	0.599	(0.011)	6.299	[0.278]	5
Inst2	(-1)	0.972	(0.001)	0.555	(0.007)	22.382	[0.000]	3
	(-2)	0.969	(0.002)	0.532	(0.006)	14.259	[0.047]	7
	(-3)	0.971	(0.001)	0.528	(0.004)	26.764	[0.005]	11
Inst3	(-1)	0.974	(0.001)	0.553	(0.008)	9.011	[0.029]	3
	(-2)	0.968	(0.002)	0.534	(0.007)	14.331	[0.046]	7
	(-3)	0.969	(0.001)	0.537	(0.005)	15.998	[0.141]	11
Inst4	(-1)	0.976	(0.005)	0.591	(0.033)	4.234	[0.237]	3
	(-2)	0.975	(0.004)	0.611	(0.018)	9.487	[0.220]	7
	(-3)	0.963	(0.002)	0.570	(0.009)	20.845	[0.035]	11
Inst5	(-1)	0.968	(0.001)	0.544	(0.007)	8.131	[0.149]	5
	(-2)	0.974	(0.001)	0.538	(0.005)	26.892	[0.005]	11
	(-3)	0.966	(0.001)	0.549	(0.005)	32.738	[0.012]	17
Panel B: $\theta_d = -0.55$, $\theta_m = -0.23$								
Inst1	(-1)	0.980	(0.009)	0.462	(0.125)	3.481	[0.062]	1
	(-2)	0.969	(0.005)	0.730	(0.244)	2.460	[0.482]	3
	(-3)	0.972	(0.005)	0.431	(0.045)	6.563	[0.255]	5
Inst2	(-1)	0.960	(0.005)	0.314	(0.023)	8.538	[0.036]	3
	(-2)	0.956	(0.005)	0.312	(0.018)	15.736	[0.028]	7
	(-3)	0.958	(0.004)	0.326	(0.025)	22.614	[0.020]	11
Inst3	(-1)	0.961	(0.006)	0.315	(0.023)	8.068	[0.045]	3
	(-2)	0.962	(0.005)	0.339	(0.022)	12.880	[0.075]	7
	(-3)	0.967	(0.004)	0.330	(0.023)	17.201	[0.102]	11
Inst4	(-1)	0.977	(0.007)	0.508	(0.191)	4.584	[0.205]	3
	(-2)	0.973	(0.004)	0.455	(0.069)	11.925	[0.103]	7
	(-3)	0.968	(0.003)	0.360	(0.019)	15.799	[0.149]	11
Inst5	(-1)	0.955	(0.006)	0.289	(0.024)	9.246	[0.100]	5
	(-2)	0.959	(0.004)	0.382	(0.029)	18.396	[0.073]	11
	(-3)	0.953	(0.004)	0.271	(0.009)	26.791	[0.061]	17

(Continued on next page)

Instruments	Lag	β	s.e.	σ	s.e.	J_T	p-value	df
Panel C: $\theta_d = -0.53$, $\theta_m = -0.36$								
Inst1	(-1)	0.976	(0.009)	0.509	(0.298)	3.808	[0.051]	1
	(-2)	0.969	(0.005)	0.788	(0.305)	2.262	[0.520]	3
	(-3)	0.973	(0.007)	0.397	(0.060)	3.481	[0.626]	5
Inst2	(-1)	0.960	(0.014)	0.260	(0.050)	7.626	[0.054]	3
	(-2)	0.966	(0.007)	0.381	(0.082)	10.347	[0.170]	7
	(-3)	0.981	(0.004)	0.479	(0.062)	17.146	[0.104]	11
Inst3	(-1)	0.967	(0.015)	0.265	(0.043)	6.000	[0.112]	3
	(-2)	0.958	(0.008)	0.294	(0.040)	11.761	[0.109]	7
	(-3)	0.974	(0.004)	0.408	(0.050)	15.640	[0.155]	11
Inst4	(-1)	0.981	(0.011)	0.826	(0.728)	5.089	[0.165]	3
	(-2)	0.973	(0.005)	0.492	(0.095)	15.579	[0.029]	7
	(-3)	0.984	(0.005)	0.428	(0.061)	28.731	[0.002]	11
Inst5	(-1)	0.966	(0.010)	0.282	(0.051)	6.873	[0.230]	5
	(-2)	0.966	(0.005)	0.458	(0.060)	13.992	[0.233]	11
	(-3)	0.979	(0.004)	0.398	(0.034)	30.744	[0.021]	17

Notes: The first column denotes types of instrument sets, which are defined as follows:
Inst1=$(const., R_{dt+1}, R_{mt+1})$,
Inst2=$(const., C_{dt+1}/C_{dt}, C_{mt+1}/C_{mt}, R_{dt+1}, R_{mt+1})$,
Inst3=$(const., C_{dt+1}/C_{dt}, C_{mt+1}/C_{mt}, P_{dt+1}/P_{dt}, P_{mt+1}/P_{mt})$,
Inst4=$(const., P_{dt+1}/P_{dt}, P_{mt+1}/P_{mt}, R_{dt+1}, R_{mt+1},)$,
Inst5=$(const., C_{dt+1}/C_{dt}, C_{mt+1}/C_{mt}, P_{dt+1}/P_{dt}, P_{mt+1}/P_{mt}, R_{dt+1}, R_{mt+1})$,
where $R_{it+1} = (1 + r_{t+1})P_{it}/P_{it+1}$ for $i = d, m$. J_T denotes Hansen's (1982) J test of the overidentifying restrictions, which has an asymptotic χ^2 distribution with degrees of freedom as reported in the final column (df).

Table 5
Stationarity Tests for Growth Rates of the Estimated Composite Good

(θ_d, θ_m)	G(0,1)	G(0,2)	G(0,3)
(-0.74,-0.42)	0.001	0.036	0.239
	[0.978]	[0.982]	[0.971]
(-0.55,-0.23)	0.461	0.466	0.832
	[0.497]	[0.792]	[0.842]
(-0.53,-0.36)	0.416	0.431	0.822
	[0.519]	[0.806]	[0.844]

Notes: The growth rates of the composite good are calculated. The value of the scale factor used is $k = 1.080$. The estimated curvature parameters used are $\alpha = 3.702$ and $\gamma = 1.326$, which are taken from Panel A of Table 3. Park's (1990) G(0,q) is a χ^2 test statistic for the null hypothesis of level stationarity. P-values are in square brackets.

Figure 1
Domestic and Imported Nondurables Expenditures and Relative Price

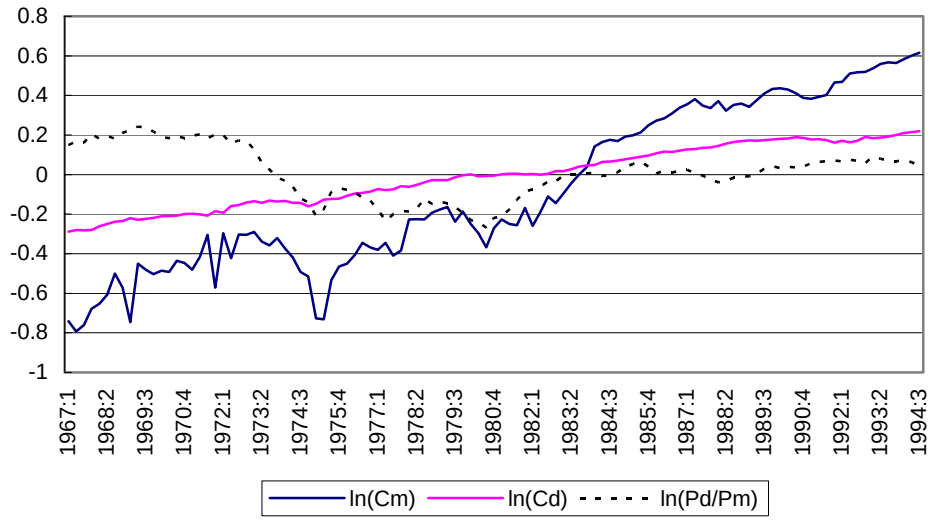

Figure 2
Growth Rates of the Composite Good