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by

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Abstract

This paper empirically reexamines the role of intertemporal substitution in United States' import demand using a model with non-separable preferences in imported and domestic goods. In previous work assuming separable preferences, a cointegration approach was widely applied for estimating preference parameters. However, this approach cannot always be used for a model with non-separable preferences because some of the available ways of introducing a stationary error into a regression make GMM estimators of parameters estimated using Euler equations inconsistent. This paper applies a simple solution to this problem and presents evidence suggesting the importance of allowing for the terms of trade in intertemporal substitution.

Key words: Intratemporal elasticity of substitution, Intertemporal elasticity of substitution, Cointegration, Generalized Method of Moments
JEL classification numbers: C22, E21, F41

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1 Introduction

In an open economy analysis, the real exchange rate, the real interest rate, and the terms of trade affect the optimal allocation of consumption, thereby influencing the adjustment of the economy to economic shocks, with consequent policy implications. On the basis of such a common understanding, most modern international macroeconomic models have been constructed to allow 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, for example, Obstfeld and Rogoff, 1996).

In these models, the way in which changes in each relative price affect households' consumption and saving decisions is governed by two parameters — the intertemporal elasticity of substitution and the intratemporal elasticity of substitution between imported and domestic goods.[1] However, much of the empirical literature on the estimation of import demand assumes an additively separable utility function and tends not to explicitly distinguish between these two elasticities (see, for example, Ceglowski 1991, Clarida 1994, 1996, Amano and Wirjanto 1996, and de la Croix and Urbain 1998). Therefore, further empirical investigations are needed in order to better understand the relationship between terms of trade shocks and household saving (i.e., intertemporal substitution).

The use of the separability assumption has obvious advantages as it allows us to focus only on the intraperiod first-order condition in the empirical specification. It also allows us to exploit a cointegration approach in estimating the intertemporal elasticity of substitution. However, when the two elasticities are distinguished by assuming the CES utility function,

[1] For a more detailed explanation, see, for example, Ostry and Reinhart (1992). In addition, an excellent description of the importance of measuring intertemporal substitution can be found in Hall (1988).

we need to take a further step and estimate the Euler equation. In this case, alternative estimation procedures are required because a stationary error introduced into the first-step regression (used in the cointegration approach) may make estimation of the second-step Euler equation inconsistent. Depending on the theoretical framework used in the analysis, there are at least two solutions to this problem. One is to introduce time non-separability of consumption by assuming either durability or habit formation, as in Ogaki and Reinhart (1998). Another solution is to exploit a forecast error arising from changing the date for one of the two goods, which is applicable to models assuming time separability.[2] Although empirical studies using these two-step estimation procedures have clarified the importance of allowing for non-separable preferences in intertemporal substitution, so far there have been few attempts to apply them in open economy analysis.

The purpose of this paper is to estimate the intratemporal and intertemporal elasticities of substitution by applying the two-step estimation procedure to a model with non-separable preferences in imported and domestic goods, and to investigate the effects of terms of trade shocks on intertemporal substitution. As pointed out in previous studies, habit formation may be a further factor that should be considered in the analysis. However, it is difficult to incorporate all of these issues into the model simultaneously. For this reason, we concentrate on non-separability across the two goods, but we consider an issue relevant to time non-separability of consumption in the empirical specification, as in Clarida (1994) and Amano and Wirjanto (1996).

In our estimation using U.S. data, we find that estimates of the intertemporal elasticity

[2] This method was originally proposed by Nishiyama (2002) and is developed by Okubo (2003) as an approach to the present problem.

of substitution are around 0.3-0.6. We interpret this as evidence in favor of more recent works, such as those by Beaudry and van Wincoop (1996), Ogaki and Reinhart (1998), and Vissing-Jørgensen (2002), which argue that the intertemporal elasticity of substitution takes positive and significant values. Moreover, our results indicate that higher substitution effects between imported and domestic goods (i.e., situations that are more sensitive to terms of trade shocks) lead to a greater intertemporal elasticity of substitution.

This paper is organized as follows. Section 2 describes the theoretical model and shows how the cointegration relationship is derived. Section 3 explains our two-step estimation procedure. Section 4 establishes the time-series properties of the data used in the analysis, and Section 5 presents the estimation results. Section 6 contains concluding remarks.

2 The Model

Suppose that a representative consumer maximizes expected lifetime utility

$$E_0\left[\sum_{t=0}^{\infty}\beta^t U(C_{dt}, C_{mt})\right], \tag{1}$$

subject to the following budget constraint:

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

where $E_t[\cdot]$ is the expectation operator conditional on the information available at time t, β is the subjective discount factor, C_{dt} is the consumption of domestic non-durable goods at time t, C_{mt} is the consumption of imported non-durable goods at time t, A_t is the stock of assets at time t, r_t is the rate of return on assets held between $t-1$ and t, Y_t is the labor income of the representative consumer at time t, and P_{it} is the price of goods i ($i = d, m$) at time t. It is assumed that the period utility function takes the CES form:

$$U(C_{dt}, C_{mt}) = \left(\frac{\sigma}{\sigma - 1}\right) [aC_{dt}^{1-(1/\epsilon)} + C_{mt}^{1-(1/\epsilon)}]^{\frac{1-(1/\sigma)}{1-(1/\epsilon)}}, \tag{3}$$

where σ is the intertemporal elasticity of substitution, ϵ is the intratemporal elasticity of substitution, and a is the parameter that denotes the weight attached to domestic consumption goods.

The optimal sequence $\{C_{dt}, C_{mt}, A_t\}$ chosen by the consumer should then satisfy the intraperiod first-order condition,

$$\frac{P_{mt}}{P_{dt}} = \frac{\partial U/\partial C_{mt}}{\partial U/\partial C_{dt}} = \frac{1}{a}\left(\frac{C_{mt}}{C_{dt}}\right)^{-\frac{1}{\epsilon}}, \tag{4}$$

and the two Euler equations,

$$E_t\left[\beta\left(\frac{\partial U/\partial C_{dt+1}}{\partial U/\partial C_{dt}}\right)(1 + r_{t+1})\frac{P_{dt}}{P_{dt+1}}\right] = 1, \tag{5}$$

$$E_t\left[\beta\left(\frac{\partial U/\partial C_{mt+1}}{\partial U/\partial C_{mt}}\right)(1 + r_{t+1})\frac{P_{mt}}{P_{mt+1}}\right] = 1, \tag{6}$$

where

$$\partial U/\partial C_{dt} = aC_{dt}^{-1/\epsilon}[aC_{dt}^{1-(1/\epsilon)} + C_{mt}^{1-(1/\epsilon)}]^{\frac{\sigma-\epsilon}{\sigma(\epsilon-1)}}, \tag{7}$$

$$\partial U/\partial C_{mt} = C_{mt}^{-1/\epsilon}[aC_{dt}^{1-(1/\epsilon)} + C_{mt}^{1-(1/\epsilon)}]^{\frac{\sigma-\epsilon}{\sigma(\epsilon-1)}}. \tag{8}$$

Since our main concern is with empirical analysis, we make the additional assumptions that $\ln C_{dt}$, $\ln C_{mt}$, and $\ln(P_{mt}/P_{ct})$ are difference stationary, as demonstrated in previous empirical work on the import demand function. Then, we may expect that the basis for estimating the parameter ϵ can be obtained by, for example, allowing for measurement errors in the variables and taking logarithms of both sides of equation (4). However, since GMM estimation presumes that there are no measurement errors, introducing such an additional error term in the model does not allow us to use the above Euler equations as the basis for estimating the remaining intertemporal parameters, β and σ. This tradeoff is a typical problem in the two-good Euler equation model.

Solving equation (4) for $C_{dt}^{-1/\epsilon}$ and $C_{mt}^{-1/\epsilon}$ and substituting the results into equations (5) and (6), respectively, we obtain the following conditions:

$$E_t\left[\beta a\left(\frac{C_{dt+1}}{C_{mt}}\right)^{-\frac{1}{\epsilon}}\left(\frac{aC_{dt+1}^{1-(1/\epsilon)}+C_{mt+1}^{1-(1/\epsilon)}}{aC_{dt}^{1-(1/\epsilon)}+C_{mt}^{1-(1/\epsilon)}}\right)^{\gamma}(1+r_{t+1})\frac{P_{mt}}{P_{dt+1}}\right]=1, \qquad (9)$$

$$E_t\left[\left(\frac{\beta}{a}\right)\left(\frac{C_{mt+1}}{C_{dt}}\right)^{-\frac{1}{\epsilon}}\left(\frac{aC_{dt+1}^{1-(1/\epsilon)}+C_{mt+1}^{1-(1/\epsilon)}}{aC_{dt}^{1-(1/\epsilon)}+C_{mt}^{1-(1/\epsilon)}}\right)^{\gamma}(1+r_{t+1})\frac{P_{dt}}{P_{mt+1}}\right]=1, \qquad (10)$$

or equivalently,

$$\beta a\left(\frac{C_{dt+1}}{C_{mt}}\right)^{-\frac{1}{\epsilon}}\left(\frac{aC_{dt+1}^{1-(1/\epsilon)}+C_{mt+1}^{1-(1/\epsilon)}}{aC_{dt}^{1-(1/\epsilon)}+C_{mt}^{1-(1/\epsilon)}}\right)^{\gamma}(1+r_{t+1})\frac{P_{mt}}{P_{dt+1}}=1+u_{1t+1}, \qquad (11)$$

$$\left(\frac{\beta}{a}\right)\left(\frac{C_{mt+1}}{C_{dt}}\right)^{-\frac{1}{\epsilon}}\left(\frac{aC_{dt+1}^{1-(1/\epsilon)}+C_{mt+1}^{1-(1/\epsilon)}}{aC_{dt}^{1-(1/\epsilon)}+C_{mt}^{1-(1/\epsilon)}}\right)^{\gamma}(1+r_{t+1})\frac{P_{dt}}{P_{mt+1}}=1+u_{2t+1}, \qquad (12)$$

where $\gamma=(\sigma-\epsilon)/\sigma(\epsilon-1)$ and u_{it+1} is a forecast error such that $E_t(u_{it+1})=0$ for $i=1,2$. Equation (9) ((10)) states that the marginal utility cost of giving up one unit of imported goods (domestic goods) at time t is equal to the expected utility gain from alternatively consuming one unit of domestic goods (imported goods) at time $t+1$. This can be interpreted as the Euler equation that determines the optimal allocation of consumption for two different goods in two consecutive times, which may be called the *cross* intertemporal substitution. It should be noted that the error terms are stationary random variables, if the gross return, $(1+r_{t+1})$, and the growth rate of marginal utility can be regarded as stationary processes (see Nishiyama, 2002 for more detailed discussions under the separability assumption). An important point here is that when this type of the model with CES utility is used, the explicit allowance for the cross-intertemporal substitution involves a redefinition of the elasticity of substitution, which is explained below.

We take logarithms of both sides of equations (11) and (12), to derive the following relations:

$$\ln(\beta a) - \frac{1}{\epsilon}\ln\left(\frac{C_{dt+1}}{C_{mt}}\right) + \gamma\ln\left(\frac{aC_{dt+1}^{1-(1/\epsilon)} + C_{mt+1}^{1-(1/\epsilon)}}{aC_{dt}^{1-(1/\epsilon)} + C_{mt}^{1-(1/\epsilon)}}\right) + \ln\left[(1+r_{t+1})\frac{P_{mt}}{P_{dt+1}}\right] = v_{1t+1}, \quad (13)$$

$$\ln\left(\frac{\beta}{a}\right) - \frac{1}{\epsilon}\ln\left(\frac{C_{mt+1}}{C_{dt}}\right) + \gamma\ln\left(\frac{aC_{dt+1}^{1-(1/\epsilon)} + C_{mt+1}^{1-(1/\epsilon)}}{aC_{dt}^{1-(1/\epsilon)} + C_{mt}^{1-(1/\epsilon)}}\right) + \ln\left[(1+r_{t+1})\frac{P_{dt}}{P_{mt+1}}\right] = v_{2t+1}, \quad (14)$$

where $v_{it+1} = \ln(1+u_{it})$ for $i = 1, 2$. Furthermore, to eliminate the unobservable term on the above equations, subtracting equation (14) from equation (13) yields

$$\left[\ln\left(\frac{C_{dt+1}}{C_{mt}}\right) - \ln\left(\frac{C_{mt+1}}{C_{dt}}\right)\right] = \theta + \epsilon\left[\ln\left(\frac{P_{mt}}{P_{dt+1}}\right) - \ln\left(\frac{P_{dt}}{P_{mt+1}}\right)\right] + e_{t+1}, \quad (15)$$

where $\theta = \epsilon[\ln(\beta a) - \ln(\beta/a)]$ and e_{t+1} is a stationary error term. The price ratio on the right-hand side of equation (15), P_{mt}/P_{dt+1}, is the price that determines the consumption expenditures on domestic goods at time $t+1$ in terms of imported goods at time t; P_{dt}/P_{mt+1} is the analogous price governing the choice between imported goods at time $t+1$ and domestic goods at time t. Thus, equation (15) can be interpreted as a definition of the elasticity of substitution that is incorporating the intertemporal structure into the usual one. In this paper, we identify the intratemporal elasticity, ϵ, based on equation (15), instead of directly using the intraperiod first-order condition as in the standard cointegration approach.

Finally, given the assumption of difference stationarity introduced previously, it is easy to verify that the two variables, $[\ln(C_{dt+1}/C_{mt}) - \ln(C_{mt+1}/C_{dt})]$ and $[\ln(P_{mt}/P_{dt+1}) -$

$\ln(P_{dt}/P_{mt+1})]$, are also difference stationary. Hence, our model can imply that these variables are cointegrated with a cointegrating vector $(1, -\epsilon)'$.

3 The Estimation Procedure

One important advantage of the above approach is that, while we have obtained equation (15) without assuming any artificial error term, the process for deriving it does not affect the specification of the Euler equation (5) or (6). Therefore, we can apply a two-step estimation procedure that combines a cointegration method with GMM, similar to the procedure followed by Ogaki and Reinhart (1998).

As mentioned above, for our model to imply the cointegration relationship, it is required that the gross return and the growth rate of marginal utility are stationary. The former is a widely recognized fact and the latter is also empirically confirmed by Ogaki and Reinhart (1998) and Okubo (2003). Thus, if the two variables in equation (15) are difference stationary, we can deal with equation (15) as the cointegrating regression. Then, applying the appropriate estimation to (15) yields a super-consistent estimate of ϵ. Unfortunately, in this step, we cannot identify an estimate of the parameter a. However, it is possible to calculate a plausible value of a from equation (4). Specifically, we utilize

$$a = \exp\left[\frac{1}{\epsilon}\ln\left(\frac{C_{dt}}{C_{mt}}\right) - \ln\left(\frac{P_{mt}}{P_{dt}}\right)\right], \tag{16}$$

where ϵ is fixed at the estimate from equation (15). We obtain the value of a by taking the exponential of the sample mean of the variable in the square brackets.

Given that equation (15) holds, since equations (5) and (6) are not independent, it is sufficient to consider equation (5) alone in the estimation of the system consisting of equations

(4)-(6). In the second step of the procedure, we apply GMM to the Euler equation (5), into which the super-consistent estimate from the first-step cointegrating regression was plugged. This approach does not alter the asymptotic properties of the GMM estimator, and it makes it possible to control for the effect of ignoring the intratemporal substitution between imported and domestic goods in estimating the intertemporal elasticity of substitution.

4 Data

We use U.S. quarterly data on consumption expenditures of imported and domestic non-durables, the corresponding price indices, and nominal interest rates (the 90-day t-bill rate), which cover the period from 1967:1 to 1994:3 (111 observations). Our measure of non-durable consumption is food plus non-oil non-durables. The 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 they all are constructed following the definition of Ceglowski (1991). Further details about the source and the construction can be found in Section II of Ceglowski's paper. 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 $\ln C_{dt}$, $\ln C_{mt}$, and $\ln(P_{mt}/P_{dt})$. Thus, in this paper, we begin by testing whether $[\ln(C_{dt+1}/C_{mt}) - \ln(C_{mt+1}/C_{dt})]$ and $[\ln(P_{mt}/P_{dt+1}) - \ln(P_{dt}/P_{mt+1})]$ are difference stationary.

Before turning to the formal tests, it will be useful to examine the changes in the two series over time. Their plots for our sample period are depicted in Figure 1. We see from the figure that while $[\ln(C_{dt+1}/C_{mt}) - \ln(C_{mt+1}/C_{dt})]$ has an explicit downward trend, this does not appear to be the case for $[\ln(P_{mt}/P_{dt+1}) - \ln(P_{dt}/P_{mt+1})]$. In other words, this difference suggests that it is appropriate to design tests of the hypothesis of difference stationarity to take trend stationarity and level stationarity, respectively, as the alternative hypothesis.[3]

We use both a modified version of the Phillips and Perron (1988) Z_α test and the Elliott et al. (1996) DF-GLS test and a modified feasible point optimal test, proposed by Ng and Perron (2001). 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. Table 1 presents the test statistics, using the lag orders selected by the modified AIC (MAIC). As can be seen from the table, the tests fail to reject the null hypothesis of difference stationarity for both variables, even at the 10% significance level. Thus, our assumptions required to derive the cointegration relationship are supported empirically.

5 Results

Next, we examine the empirical validity of equation (15), which summarizes the long-run restriction on the deterministic and stochastic trends of $[\ln(C_{dt+1}/C_{mt}) - \ln(C_{mt+1}/C_{dt})]$ and $[\ln(P_{mt}/P_{dt+1}) - \ln(P_{dt}/P_{mt+1})]$. When these series can be modeled as being I(1) processes with drift, equation (15) means that the cointegrating vector $(1, -\epsilon)'$ eliminates the deterministic trend as well as the stochastic trend. However, our previous findings indicate that the two series are both I(1) processes, but that one of them, particularly

[3] Since the model does not tell us whether the variables are integrated of order 1 (I(1)) with or without drift, the type of alternative hypothesis used may have to depend on graphical investigations. One reason for this is that the implementation of unit root tests against both level stationarity and trend stationarity does not always make it possible to identify the difference between the two. Indeed, in our unit root tests, we encountered this difficulty. In general, this may not be a crucial problem. However, in the cointegration approach, it is important because whether the variables have a drift term affects the type of cointegration to be tested (i.e., stochastic or deterministic cointegration). In the following analysis, in which our procedure assumes that both variables are I(1) with drift, the null hypothesis of deterministic cointegration was rejected. However, as later discussed in footnote 5, such modeling does not seem to be consistent with previous work in this field. Thus, we specified the alternative hypothesis following the graphical suggestion.

$[\ln(P_{mt}/P_{dt+1}) - \ln(P_{dt}/P_{mt+1})]$, may not have drift. In this case, our model does not involve the restriction that the cointegrating vector eliminates the deterministic trend[4] ; therefore, we allow for a trend in the cointegration relationship.[5]

Table 2 reports the estimation results of equation (15) with a trend. For estimating the equation, we use Phillips and Hansen's (1990) fully modified (FM) OLS and Park's (1992) canonical cointegrating regression (CCR). In the CCR procedure, Park's (1990) H(p,q) test is also used to test the null hypothesis of cointegration.

The point estimates of ϵ have the theoretically expected sign and are significantly estimated at a range of 0.626 to 0.628. This magnitude of ϵ may appear small, but it is close to that obtained by the previous studies (for example, Ceglowski, 1991, Amano and Wirjanto, 1996, and Nishiyama, 2002) using the model with the assumption of $\sigma = \epsilon$. The values of the H(1,2), H(1,3), and H(1,4) test statistics were 0.002 (p-value=0.969), 1.423 (p-value=0.491), and 2.336 (p-value=0.506), respectively. Thus, we do not reject the null hypothesis of stochastic cointegration even at the 10% significance level.

Table 3 reports the estimation results of equation (5) based on the second step of GMM, where ϵ and a are set at the CCR estimate and the value obtained using (16). The instrument variables used in this paper consist of a constant, the real gross returns, the growth rate of C_{dt},

[4] To see this, let y_t be an I(1) variable with drift and let x_t be an I(1) variable without drift. In this case, a linear combination by a cointegrating vector $(1, -\epsilon)'$ is as follows: $y_t - \epsilon x_t = \mu_y t + (\sum_{s=1}^{t} u_{ys} - \epsilon \sum_{s=1}^{t} u_{xs})$, where μ_y is a drift term of y_t and u_{yt} and u_{xt} are stationary variables. Thus, it follows that there is no particular relation between the deterministic trend and the cointegrating vector.

[5] An empirical specification with a trend term was also adopted by Clarida (1994) and Amano and Wirjanto (1996) to capture the effects of quality improvements, which are difficult to incorporate into the two-good model. Ideally, it is desirable to include a trend in the intraperiod first-order condition in the model. One conceivable way to do this may be to introduce an exponential deterministic trend that embodies the transformation of purchases into services, as in Ogaki and Park (1998) and de la Croix and Urbain (1998). However, this solution makes the second step of GMM difficult because the existence of such a trend and time non-separability complicates the Euler equation, even if the first step of cointegration is applicable. Therefore, for the purpose of comparison with the previous studies, we chose to introduce the trend term in the regression.

the growth rate of C_{mt}, and the growth rate of C_{dt}/C_{mt}. To control for the time aggregation problem by the use of quarterly data, the instrument variables are all lagged two periods. Moreover, we calculate the weighting matrix using the VAR(1) prewhitening technique of Andrews and Monahan (1992) under the assumption that the disturbance follows an MA(1) process with an unknown coefficient. Each GMM estimation is iterated five times.

In Panel A of Table 3, the point estimates of the intertemporal elasticity of substitution, σ, are positive and significantly estimated around 0.3-0.5. Ogaki and Reinhart (1998), who estimate the intertemporal elasticity σ using U.S. quarterly data, find that this parameter is around 0.4, based on a similar CES utility function. de la Croix and Urbain (1998) also report that the estimates of σ for imported and domestic goods seem to be around 0.3, based on a comprehensive survey. Therefore, our estimates fall into a reasonable range within our reading of the literature in this field.[6] In addition, Hansen's J-test statistic does not reject the model even at the 10% significance level. Overall, the above results show that our model is supported statistically.

On the other hand, we find that the rejection of the separability assumption ($\sigma = \epsilon$) is somewhat weak. Given the strong rejection of $\sigma = \epsilon$ for developing countries (see, for example, Ostry and Reinhart, 1992 and Ogaki, Ostry, and Reinhart, 1996), our estimation results may be interpreted as evidence indicating that the difference between σ and ϵ, that is, the impact of changes in the terms of trade on consumption choices, is smaller for developed countries such as the United States.

[6] We also conducted the base runs without adding the constant premium of 2%. The resulting estimates of σ were 0.495 (standard error=0.074), 0.433 (standard error=0.085), and 0.350 (standard error=0.110), respectively. However, β was estimated to be greater than one. Although the estimates of β are somewhat puzzling, de la Croix and Urbain (1998) and Nishiyama (2002) also find similar estimates using different models but the same data set. Thus, we may be able to state that this puzzling result is not a problem unique to our CES specification.

To further examine the relation between ϵ and σ, Panel B of Table 3 reports the results when the value of ϵ was increased by two standard errors. This experiment corresponds to the case in which consumers have stronger incentives to substitute between imported and domestic goods. The results indicate that the increase in the value of ϵ leads to a rise in σ.[7] The J-test statistic still does not reject the model under such a change in ϵ. This finding supports our view that allowing for terms of trade shocks is also important in evaluating intertemporal substitution.

6 Conclusion

This paper has applied an alternative estimation procedure for estimating preference parameters, and has attempted to estimate the intratemporal and intertemporal elasticities of substitution using a model with non-separable preferences in imported and domestic goods. An advantage of our procedure is that a cointegration approach can be applied even when the non-separability of preferences is assumed. This property makes it possible to distinguish the difference in the magnitude of the two elasticities explicitly and to examine empirically the effects of the terms of trade in intertemporal substitution.

The estimation results obtained in this paper indicate that the intertemporal elasticity of substitution is significantly different from zero and takes values around 0.3-0.6. In addition, our results reveal that the degree of *intra*temporal substitution between imported and domestic goods affects the magnitude of the intertemporal elasticity of substitution. This finding suggests that the terms of trade is an important factor in the measurement of in-

[7] Again, in the case without adding the constant premium, the results for σ were highly similar to those reported in Panel B. Thus, we can safely say that our results do not depend on whether the constant premium is added to the real return.

tertemporal substitution, in addition to the existence of time non-separability, as claimed by some previous studies.

However, in an empirical study of an open economy, the relative importance of these factors will be different depending on the economic environment or the international position of the countries analyzed. Thus, further investigation based on other countries' data is of interest for future research.

On the other hand, like the previous studies on the estimation of import demand, we encountered at least two difficulties in estimating preference parameters: the capture of the effects of quality improvements and the point estimate of the discount factor being greater than one. As already mentioned in the text, these problems may be resolved by extending the model to allow for an exponential deterministic trend. However, this remedy makes the second step of GMM difficult and complicated. This problem is expected to be resolved in the future, but our two-step approach will remain useful for estimating such a model with non-separable preferences.

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Table 1
Tests for the Null Hypothesis of Difference Stationarity

Variable	k_{maic}	$\bar{M}Z_{\alpha}^{GLS}$	DF^{GLS}	$\bar{M}P_{T}^{GLS}$
$\ln(C_{dt+1}/C_{mt}) - \ln(C_{mt+1}/C_{dt})$	4	-9.977	-2.132	9.178
$\ln(P_{mt}/P_{dt+1}) - \ln(P_{dt}/P_{mt+1})$	6	-2.430	-1.068	10.010

Notes: k_{maic} denotes the number of lags determined by the modified AIC, where an upper bound on the lag length is set at the integer part of $12(T/100)^{1/4}$. Critical values of the $\bar{M}Z_{\alpha}^{GLS}$, DF^{GLS}, and $\bar{M}P_{T}^{GLS}$ tests are taken from Table I of Ng and Perron (2001). The test regressions include a trend term for $\ln(C_{dt+1}/C_{mt}) - \ln(C_{mt+1}/C_{dt})$, whereas they only have a constant term for $\ln(P_{mt}/P_{dt+1}) - \ln(P_{dt}/P_{mt+1})$. The factors $\bar{c}$ required to construct the GLS detrended series are set at -13.5 and -7.0, respectively, following Elliott et al. (1996).

Table 2
Estimation Results of the Cointegrating Regression

Parameters	CCR	FM
trend	-0.016	-0.016
	(0.001)	(0.001)
θ	1.042	1.036
	(0.056)	(0.057)
ϵ	0.626	0.628
	(0.098)	(0.108)

Notes: CCR denotes Park's (1992) canonical cointegrating regression and FM denotes Phillips and Hansen's (1990) fully modified OLS. The CCR and FM estimates are based on the quadratic spectral kernel and the VAR(1) prewhitening technique of Andrews and Monahan (1992). Standard errors are in parentheses.

Table 3
Generalized Method of Moments Results

Instruments	ϵ	a	β	σ	J_T	df
Panel A: Base Run Results						
Inst1	0.626	1.139	0.985	0.496	4.550	4
			(0.003)	(0.074)	[0.337]	
Inst2	0.626	1.139	0.987	0.434	1.961	3
			(0.004)	(0.085)	[0.581]	
Inst3	0.626	1.139	0.993	0.352	0.010	2
			(0.009)	(0.110)	[0.995]	
Panel B: The Case of Increased Substitution Effects						
Inst1	0.822	1.139	0.984	0.609	3.757	4
			(0.002)	(0.100)	[0.440]	
Inst2	0.822	1.139	0.987	0.533	2.017	3
			(0.004)	(0.119)	[0.569]	
Inst3	0.822	1.139	0.991	0.415	0.001	2
			(0.008)	(0.152)	[0.999]	

Notes: Inst1 $= \{1, C_{dt-1}/C_{dt-2}, C_{mt-1}/C_{mt-2}, C_{dt-2}/C_{mt-2}, R_{dt-1}, R_{mt-1}\}$, Inst2 $=$ Inst1 excluding R_{mt-1}, and Inst3 $=$ Inst1 excluding R_{mt-1} and C_{dt-2}/C_{mt-2}, where R_d and R_m denote the real gross returns defined using the domestic goods deflator and the imported goods deflator, respectively. J_T denotes Hansen's (1982) J-test of the overidentifying restrictions, which has asymptotic χ^2 distributions with the degrees of freedom reported in the last column (df). Standard errors are in parentheses and p-values are in square brackets. The value of ϵ is from Table 2 and that of a is based on equation (16).

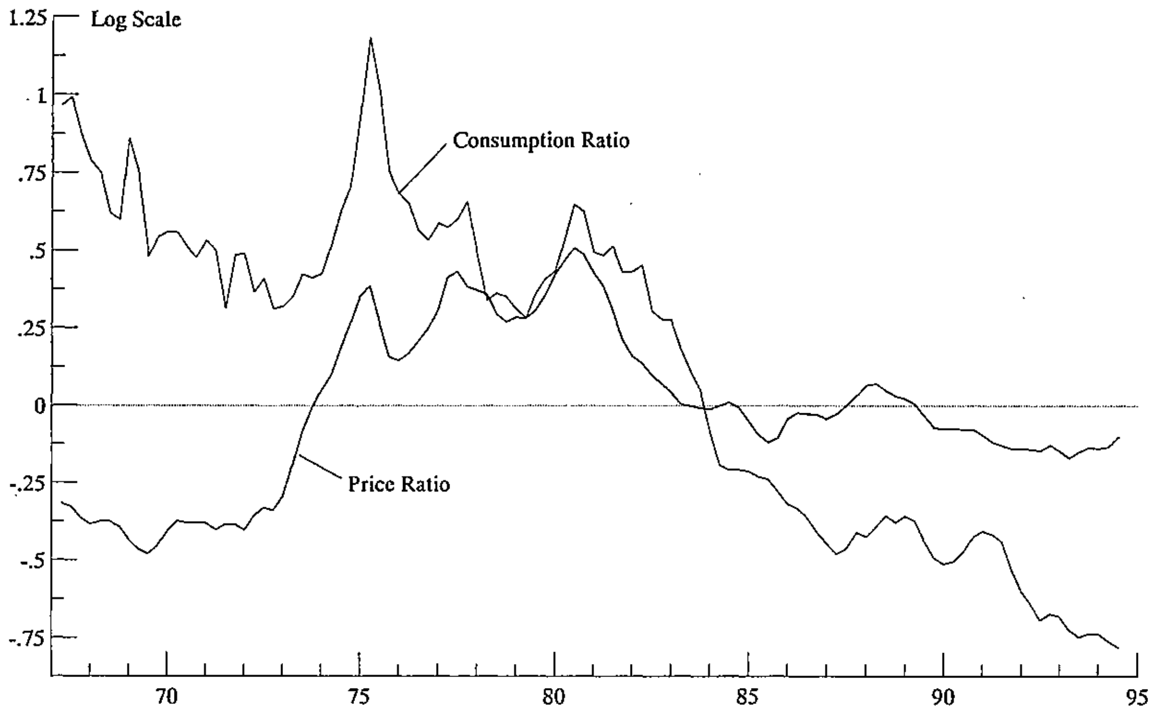

Figure 1: Consumption Ratio and Price Ratio 1967Q2-1994Q3