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Intertemporal Substitution and Nonhomothetic Preferences [*]

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Abstract

This study uses a model with nonseparable and nonhomothetic preferences to estimate the intertemporal elasticity of substitution (IES). We show that, while the assumption of homotheticity is strongly rejected, the estimated IES is positive and significant.

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

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1 Introduction

As Ogaki and Reinhart (1998a,b) demonstrated, incorporating nonseparability between nondurable and durable goods into a model yields a more plausible estimate of the intertemporal elasticity of substitution (IES) in consumption. This finding is important because it provides a 'theoretical' explanation for the zero and negative IES estimates obtained by Hall (1988).[1] To allow for nonseparability of preferences, Ogaki and Reinhart's model assumes constant elasticity of substitution (CES) preferences that are homothetic. However, as Deaton (1992, pp.8-9), among others, have emphasized, the homotheticity assumption contradicts virtually all household budget studies or time-series analysis of expenditure patterns.[2] Inevitably, those facts make the adoption of CES-type preferences in an empirical model of consumption doubtful.

To what extent does this deficiency of the model bias the IES estimates? Does the additional requirement of nonhomotheticity affect the IES estimates? The purpose of this note is to answer these questions by examining whether Ogaki and Reinhart's results are robust to nonhomotheticity of preferences. Specifically, we investigate whether the estimated IES is significantly different from zero under nonseparable and nonhomothetic preferences.[3]

In this note, we introduce the addilog-type utility function with nonseparability in nondurable and durable goods. This slight generalization enables one to formally test the null hypothesis of homothetic preferences. Using Ogaki and Reinhart's data, we find that the

[1] For a recent econometric explanation, see, e.g., Yogo (2004).

[2] More recent empirical studies also suggest the need to relax the assumption of homothetic preferences. For example, Ait-Sahalia et al. (2004) find that nonhomothetic utility functions for luxury and basic consumption goods make an important contribution to explaining asset-pricing puzzles.

[3] Another conceivable strategy is to compute the IES for total expenditure (i.e., budget share weighted sum of the IES for each good) along the lines of Atkeson and Ogaki (1996) and Browning and Crossley (2000). To facilitate comparison with the previous studies, we address the homotheticity issue from the aspect of estimation and testing of preference parameters, and concentrate on the IES for a composite good.

homotheticity assumption is strongly rejected. However, even when nonhomotheticity of preferences is taken into account, our point estimates for the IES are positive and significant. This suggests that the introduction of intratemporal substitution between nondurable and durable goods remains important.

2 Model and Methodology

Suppose that a representative consumer maximizes the following lifetime utility function at time 0:

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

where β is the subjective discount factor, C_{1t} is the consumption of nondurable goods in period t, S_t is the stock of durable goods C_{2t} in period t, and $E_t[\cdot]$ denotes expectations conditional on the information available at time t. It is assumed that durables begin to yield services in the same period that they are purchased: $S_t = C_{2t} + \delta C_{2t-1} + \delta^2 C_{2t-2} + \cdots$, where $1-\delta$ denotes the depreciation rate for durable goods.

We assume a period utility function for nondurables consumption and service flows of the following form:

$$u(C_{1t}, S_t) = \frac{\left[\dfrac{C_{1t}^{1-\alpha}}{1-\alpha} + \dfrac{S_t^{1-\gamma}}{1-\gamma}\right]^{1-1/\sigma}}{1-1/\sigma}, \tag{2}$$

where $\alpha > 0$ and $\gamma > 0$ are the curvature parameters and σ is the IES. Thus, we do not constrain preferences to be separable and homothetic. When $\alpha = \gamma = 1/\epsilon$, this utility function incorporates one similar to the CES-type (i.e., nonseparable but homothetic) utility function examined by Ogaki and Reinhart (1998a,b), in which ϵ is interpreted as the intratemporal elasticity of substitution. When $1/\sigma = 0$, on the other hand, it reduces to the addilog utility function familiar in the cointegration literature.

Let P_{it} be the purchase price of consumption good C_{it}, and let Q_t be the user cost of the durable good. The relative price of nondurable and durable goods is defined as $P_t \equiv P_{2t}/P_{1t}$. Since, by defining utility over the service flow from durables, one can treat durable goods within the class of time-separable preferences, the intraperiod first-order condition for an optimum is given by

$$Q_t = \frac{\partial U/\partial S_t}{\partial U/\partial C_{1t}} = \frac{S_t^{-\gamma}}{C_{1t}^{-\alpha}}, \tag{3}$$

and the Euler equation generated from the period utility function (2) can be written as

$$E_t \left[\beta R_{t+1} \left(\frac{C_{1t+1}}{C_{1t}} \right)^{-\alpha} \left[\frac{(1-\gamma)C_{1t+1}^{1-\alpha} + (1-\alpha)S_{t+1}^{1-\gamma}}{(1-\gamma)C_{1t}^{1-\alpha} + (1-\alpha)S_t^{1-\gamma}} \right]^{-1/\sigma} \right] = 1, \tag{4}$$

where R_{t+1} is the gross return on any asset in terms of the nondurable good at time $t+1$. Given $1/\sigma = 0$ (i.e., separable preferences), equation (4) is equivalent to the one-good model estimated by Ogaki and Reinhart (1998a,b). Hence, in our empirical work using Ogaki and Reinhart's data sets, we can concentrate on estimation of the two-good model.

For practical purposes, the intraperiod first-order condition (3) involves an approximation of the user cost. Therefore, we consider estimating the curvature parameters from another intraperiod first-order condition:

$$\begin{aligned} P_t &= \frac{\partial U/\partial C_{2t}}{\partial U/\partial C_{1t}} \\ &= \frac{E_t \left[\sum_{i=0}^{\infty} \beta^i \delta^i u'_{C_2}(C_{1t+i}, S_{t+i}) \right]}{u'_{C_1}(C_{1t}, S_t)}, \end{aligned} \tag{5}$$

where marginal utilities of C_{1t} and C_{2t}, respectively, are given by

$$u'_{C_1}(C_{1t}, S_t) = C_{1t}^{-\alpha} \left[\frac{C_{1t}^{1-\alpha}}{1-\alpha} + \frac{S_t^{1-\gamma}}{1-\gamma} \right]^{-1/\sigma}, \tag{6}$$

$$u'_{C_2}(C_{1t}, S_t) = S_t^{-\gamma} \left[\frac{C_{1t}^{1-\alpha}}{1-\alpha} + \frac{S_t^{1-\gamma}}{1-\gamma} \right]^{-1/\sigma}. \tag{7}$$

Multiplying both sides of equation (5) by $C_{1t}^{-\alpha}/C_{2t}^{-\gamma}$ and rearranging terms on the right-hand side of the resulting equation yields the following relation:

$$P_t \frac{C_{1t}^{-\alpha}}{C_{2t}^{-\gamma}} = E_t \left[\sum_{i=0}^{\infty} \beta^i \delta^i \frac{u'_{C_1}(C_{1t+i}, S_{t+i})}{u'_{C_1}(C_{1t}, S_t)} \left(\frac{C_{1t}}{C_{1t+i}} \right)^{-\alpha} \left(\frac{S_{t+i}}{C_{2t}} \right)^{-\gamma} \right]. \tag{8}$$

As we explain in Section 3, it is not possible to reject the null hypothesis that $\ln(C_{1t})$, $\ln(C_{2t})$, and $\ln(P_t)$ are difference stationary with drift. It follows from (8) that if the growth rate of marginal utility is stationary, $\ln(P_t) - \alpha \ln(C_{1t}) + \gamma \ln(C_{2t})$ is stationary. This follows from the fact that the right-hand side of (8) can be expressed as a function of the stationary variables, C_{1t}/C_{1t+i} and S_{t+i}/C_{2t}. Thus, the theory implies that $\ln(C_{1t})$, $\ln(C_{2t})$, and $\ln(P_t)$ are deterministically cointegrated (i.e., they are stochastically cointegrated with the deterministic cointegration restriction).

Given (6), we see that the growth rate of marginal utility is a function of the growth rate of C_{1t} and the growth rate of the composite good $[C_{1t}^{1-\alpha}/(1-\alpha) + S_t^{1-\gamma}/(1-\gamma)]$. However, the above difference stationarity assumption does not ensure the stationarity of the latter growth rate. Therefore, the growth rate of marginal utility could be nonstationary. In this study, we assume the stationarity of the growth rate of the composite good and then test the validity of this assumption by using the estimated composite good.

In our empirical work, we estimate the cointegrating regression

$$\ln(C_{2t}) = const. - \frac{1}{\gamma} \ln(P_t) + \frac{\alpha}{\gamma} \ln(C_{1t}) + u_t \tag{9}$$

by applying Park's (1992) canonical cointegrating regression (CCR) method, where u_t denotes a stationary error term. In general, it is possible to use other variables as the dependent variable. We choose $\ln(C_{2t})$ as a dependent variable to facilitate comparison with the intratemporal elasticity ϵ in Ogaki and Reinhart's model. In addition to testing deterministic cointegration by using Park's (1990) H(p,q) statistics, we test the null hypothesis of homothetic preferences (i.e., $\alpha = \gamma$) by using the K statistic, which has a χ^2 distribution with one degree of freedom under the null hypothesis. We then apply Hansen's (1982) generalized method of moments (GMM) to the Euler equation (4) with the CCR estimates of α and γ and thereby estimate the IES.

3 Empirical Results

The data used in this study are the same as those used by Ogaki and Reinhart (1998a). They are quarterly and cover the period from 1947:1 to 1983:4. Data on nondurables (excluding clothing) are from the National Income and Product Accounts (NIPA), and data on durables are from the NIPA and Gordon (1990). Per capita real consumption series are constructed by dividing these series by the total population, including armed forces overseas. Nominal interest rates are nominal after-tax rates and are defined using the three-month Treasury bill rate and Barro and Sahasakul's (1983) marginal average tax rate series. In calculating the service flow series S_t, for the depreciation rate, we use $\delta = 0.92$ and $\delta = 0.96$, together with $\delta = 0.94$, used predominantly by Ogaki and Reinhart (1998a), to check for robustness.[4] We consider two periods for estimation: 1947:2–1983:4 and 1951:1–1983:4.

As a preliminary step, we examine the descriptive and statistical properties of the data. Table 1 shows that the (budget) shares of nondurable and durable expenditure in total expenditure (defined as the sum of durable and nondurable expenditure) changed substantially over the sample period. This suggests that the data do not support the assumption of homotheticity. To confirm the trend properties of the data series, we perform Park's (1990) J(1,5) test for unit roots in $\ln(C_{1t})$, $\ln(C_{2t})$ and $\ln(P_t)$. For the period from 1947:2 to 1983:4,

[4] Values for δ of 0.92, 0.94 and 0.96 imply annual depreciation rates of about 28%, 22% and 15%, respectively.

the value of the J(1,5) test was 1.714 for $\ln(C_{1t})$. When Gordon's data were used, values of 0.690 and 2.252 for $\ln(C_{2t})$ and $\ln(P_t)$, respectively, were obtained. Using the NIPA data on durables, we obtained values of 0.948 and 7.328 for $\ln(C_{2t})$ and $\ln(P_t)$, respectively. Results for the period from 1951:1 to 1983:4 were similar. Thus, the null hypothesis of difference stationarity with drift is not rejected.

Table 2 reports the CCR results. Panel A corresponds to equation (9) and Panel B reports results from Ogaki and Reinhart's specification. In Panel A, the H(0,1) test fails to reject the deterministic cointegration restriction, even at the 10% level. The H(1,q) tests fail to reject the null hypothesis of stochastic cointegration, even at the 10% level. For both sample periods, the point estimates of $1/\gamma$ and α/γ are significantly different from zero, and have the theoretically expected signs. More importantly, for all cases, the K test strongly rejects the null hypothesis of homothetic preferences. This is consistent with the observations in Table 1.

Panel B of Table 2 (in which the first and second rows correspond to the estimation results of Ogaki and Reinhart (1998a)), shows that the cointegrating regression that imposes homotheticity, $\alpha = \gamma$, is supported for the 1951:1–1983:4 period. In other words, from the H(p,q) tests, we have evidence to support two different cointegrating regressions. However, the estimates of the preference parameters seem to differ significantly between the two.

Table 3 presents the GMM results. With β fixed at 0.990, since we were unable to estimate σ because of convergence problems, we estimated β and σ simultaneously. Hence, it should be noted that the magnitude of our IES estimates is not directly comparable with that of Ogaki and Reinhart (1998a).[5]

[5] The issue is whether our model, which incorporates nonseparable and nonhomothetic preferences, can yield non-zero or non-negative IES estimates, which is consistent with Ogaki and Reinhart's main motivation.

The first two rows of each panel of Table 3 report results based on $\delta = 0.94$. The instrumental variables are those used by Ogaki and Reinhart (1998a) plus the growth rate of durables.[6] In the following analysis, all instruments are lagged two periods to control for the effect of time aggregation, and the values of α and γ are based on the estimation results of Table 2. When Gordon's data are used, for the 1951:1–1983:4 period, Hansen's J test rejects the model at the 10% level, but not at the 5% level. On the other hand, for the sample period 1947:2–1983:4, the J test does not reject the model, even at the 10% level. Our point estimates of σ are positive and significantly different from zero. The separability assumption (i.e., $1/\sigma = 0$) is rejected despite the difference in the sample periods. Results from the NIPA durables in Panel B are similar to the above results from Gordon's data.

The third and fourth rows of each panel report results based on $\delta = 0.92$, while the fifth and sixth rows report results based on $\delta = 0.96$. For both data sets, the results are similar to those based on $\delta = 0.94$ except in the case of $\delta = 0.92$ for the 1951:1–1983:4 period. Therefore, our results are robust to the value of the depreciation rate and the choice of data sets.[7]

We should emphasize that while our model nests the CES-type utility function when $\alpha = \gamma$, it does not completely incorporate Ogaki and Reinhart's CES utility function as a special case. This probably explains why we were unable to obtain a plausible estimate of the IES when the Euler equation with $\alpha = \gamma = 0.857$ (the inverse of $1/\gamma = 1.167$, i.e., Ogaki and Reinhart's estimate) was used in the second step of GMM. For example, for $\delta = 0.94$,

[6] We included the growth rate of C_{2t} as an instrument to compensate for the reduction in the degrees of freedom due to the joint estimation of β and σ.

[7] We also tried the instrument set used by Ogaki and Reinhart (1998a). The results are similar to those reported in Table 3. For example, for $\delta = 0.94$, when Gordon's data are used, the estimates of σ are 0.157 (s.e.=0.047) and 0.201 (s.e=0.090) for the 1947:2–1983:4 and 1951:1–1983:4 periods, respectively. When the NIPA data on durables are used, the corresponding estimates are 0.151 (s.e.=0.047) and 0.143 (s.e.=0.039), respectively.

the estimate of σ was 10.000 (s.e.=552.081).[8]

However, under the separability assumption, it should be noted that our model is equivalent to Ogaki and Reinhart's model, as we explained in Section 2. That is, the one-good model yields negative IES estimates. Hence, our results convey the same message as Ogaki and Reinhart's: ignoring the intratemporal substitution effect by imposing separability when estimating the IES biases the estimates downwards.

Finally, we examine the stationarity of the growth rate of the composite good. Table 4 presents the test results for the null hypothesis of level stationarity based on Park's (1990) G(0,q) test. For the 1947:2–1983:4 period, which includes the period of restocking durable goods after World War II, the evidence in favor of stationarity is weak. That is, in our framework, the inclusion of this unusual period weakens the basis for the cointegrating regression. Nevertheless, including this period did not seem to affect our estimation and test results in Table 2, unlike Ogaki and Reinhart's results.[9] This test result for stationarity suggests that the sample period used for analysis should begin at 1951:1. However, Tables 2 and 3 indicate that, even if we focus only on the estimation results for the 1951:1–1983:4 period, the main findings of this study remain the same.

4 Conclusions

In this note, we introduced and estimated a model incorporating both nonseparability and nonhomotheticity of preferences to confirm Ogaki and Reinhart's (1998a) results. This attempt is important because Ogaki and Reinhart's model imposes homotheticity over non-

[8] When σ was greater than 10, we penalized GMM estimation by multiplying the disturbance term by $1 + (|\sigma| - 10)^2$.

[9] As Gallant and White (1988) and Andrews and McDermott (1995) showed, the assumption of stationarity, which is required for GMM, can be relaxed to some extent. Hence, this violation is not expected to substantially affect the GMM results.

durable and durable goods. We found additional evidence that supports the importance of modeling nonseparability for nondurable and durable goods. Our results indicate that the intertemporal elasticity takes a low value, but is not zero.

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Table 1
Budget Share Changes over the Sample Period

Good	1947:1	1951:1	1983:4
Panel A: NIPA nondurables+Gordon durables			
Nondurables	91.5%	86.5%	71.9%
Durables	8.5%	13.5%	28.1%
Panel B: NIPA nondurables+NIPA durables			
Nondurables	83.9%	78.5%	70.6%
Durables	16.1%	21.5%	29.4%

Note: The budget share is defined as the ratio of expenditure on each good to total expenditure on nondurable and durable goods.

Table 2
Cointegrating Regression Results

Sample Period	Durables	$1/\gamma$	α/γ	H(0,1)	H(1,2)	H(1,3)	H(1,4)	K
Panel A: Nonhomothetic Preferences								
1947:2–1983:4	Gordon	0.763	2.289	0.185	0.056	0.083	2.489	18.048
		(0.127)	(0.303)	[0.667]	[0.813]	[0.959]	[0.477]	[0.000]
1951:1–1983:4	Gordon	0.701	2.469	0.041	0.279	0.791	0.993	50.419
		(0.084)	(0.207)	[0.840]	[0.598]	[0.673]	[0.803]	[0.000]
1947:2–1983:4	NIPA	0.811	1.837	0.956	0.978	1.144	2.792	25.556
		(0.178)	(0.165)	[0.328]	[0.323]	[0.564]	[0.425]	[0.000]
1951:1–1983:4	NIPA	0.716	1.941	1.935	1.172	1.321	1.486	51.671
		(0.125)	(0.131)	[0.164]	[0.279]	[0.517]	[0.686]	[0.000]
Panel B: Homothetic Preferences								
1947:2–1983:4	Gordon	1.242		7.576	0.004	0.754	1.923	
		(0.098)		[0.006]	[0.947]	[0.686]	[0.589]	
1951:1–1983:4	Gordon	1.167		3.490	0.009	1.750	2.499	
		(0.099)		[0.062]	[0.924]	[0.417]	[0.476]	
1947:2–1983:4	NIPA	1.389		9.310	0.019	0.624	2.358	
		(0.268)		[0.002]	[0.892]	[0.732]	[0.501]	
1951:1–1983:4	NIPA	1.160		2.736	0.419	0.841	1.308	
		(0.242)		[0.098]	[0.517]	[0.657]	[0.727]	

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 3
Generalized Method of Moments Results

Sample Period	δ	α	γ	β	σ	J_T
Panel A: Gordon durables data						
1947:2–1983:4	0.94	3.001	1.311	0.946	0.136	2.011
				(0.009)	(0.034)	[0.734]
1951:1–1983:4	0.94	3.523	1.427	0.961	0.218	8.012
				(0.012)	(0.099)	[0.091]
1947:2–1983:4	0.92	3.001	1.311	0.952	0.160	3.217
				(0.009)	(0.049)	[0.522]
1951:1–1983:4	0.92	3.523	1.427	0.973	0.387	7.679
				(0.011)	(0.278)	[0.104]
1947:2–1983:4	0.96	3.001	1.311	0.947	0.154	0.851
				(0.008)	(0.033)	[0.931]
1951:1–1983:4	0.96	3.523	1.427	0.947	0.151	6.753
				(0.011)	(0.040)	[0.150]
Panel B: NIPA durables data						
1947:2–1983:4	0.94	2.265	1.233	0.960	0.148	2.509
				(0.006)	(0.045)	[0.643]
1951:1–1983:4	0.94	2.712	1.397	0.962	0.159	7.829
				(0.007)	(0.046)	[0.098]
1947:2–1983:4	0.92	2.265	1.233	0.964	0.175	3.144
				(0.006)	(0.062)	[0.534]
1951:1–1983:4	0.92	2.712	1.397	0.974	0.301	8.095
				(0.007)	(0.165)	[0.088]
1947:2–1983:4	0.96	2.265	1.233	0.958	0.154	1.398
				(0.007)	(0.040)	[0.844]
1951:1–1983:4	0.96	2.712	1.397	0.959	0.162	5.652
				(0.007)	(0.042)	[0.227]

Notes: The values of α and γ are based on the estimates in Panel A of Table 2: for the 1947:2–1983:4 period, $\gamma = 1/0.763 = 1.311$ and $\alpha = 1.311 \times 2.289 = 3.001$; $\gamma = 1/0.811 = 1.233$ and $\alpha = 1.233 \times 1.837 = 2.265$. For the 1951:1–1983:4 period, $\gamma = 1/0.701 = 1.427$ and $\alpha = 1.427 \times 2.469 = 3.523$; $\gamma = 1/0.716 = 1.397$ and $\alpha = 1.397 \times 1.941 = 2.712$. J_T denotes Hansen's (1982) J-test of the overidentifying restrictions with four degrees of freedom. Standard errors are in parentheses, and p-values are in square brackets.

Table 4
Stationarity Tests for Growth Rates of the Composite Good

Sample Period	δ	α	γ	G(0,1)	G(0,2)	G(0,3)
Panel A: Gordon durables data						
1947:2–1983:4	0.94	3.001	1.311	4.962	7.922	11.763
				[0.026]	[0.019]	[0.008]
1951:1–1983:4	0.94	3.523	1.427	2.447	2.594	4.948
				[0.118]	[0.273]	[0.176]
1947:2–1983:4	0.92	3.001	1.311	4.408	6.846	11.126
				[0.036]	[0.033]	[0.011]
1951:1–1983:4	0.92	3.523	1.427	0.821	2.005	3.227
				[0.365]	[0.367]	[0.358]
1947:2–1983:4	0.96	3.001	1.311	5.600	8.952	12.133
				[0.018]	[0.011]	[0.007]
1951:1–1983:4	0.96	3.523	1.427	5.117	5.355	8.035
				[0.024]	[0.069]	[0.045]
Panel B: NIPA durables data						
1947:2–1983:4	0.94	2.265	1.233	4.847	8.505	12.704
				[0.028]	[0.014]	[0.005]
1951:1–1983:4	0.94	2.712	1.397	2.667	2.721	7.056
				[0.102]	[0.257]	[0.070]
1947:2–1983:4	0.92	2.265	1.233	4.357	7.637	12.436
				[0.037]	[0.022]	[0.006]
1951:1–1983:4	0.92	2.712	1.397	1.008	1.280	4.699
				[0.315]	[0.527]	[0.195]
1947:2–1983:4	0.96	2.265	1.233	5.457	9.365	12.788
				[0.019]	[0.009]	[0.005]
1951:1–1983:4	0.96	2.712	1.397	5.042	5.978	9.523
				[0.025]	[0.050]	[0.023]

Notes: G(0,1), G(0,2), and G(0,3) are χ^2 test statistics for the null hypothesis of level stationarity. P-values are in square brackets.