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Ogaki, Masao and Carmen M. Reinhart

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Masao Ogaki
Ohio State University

and

Carmen M. Reinhart
International Monetary Fund

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Abstract

As pointed out by Hall (1988), intertemporal substitution by consumers is a central element of many modern macroeconomic and international models. For example, many of the policy implications of an endogenous growth model studied by Barro (1990) depends on the assumption that the intertemporal elasticity of substitution is positive. In estimating the intertemporal elasticity of substitution (IES), however, Hall (1988) finds that when time aggregation is taken into account, his point estimates are small and not significantly different from zero. Hall concludes that the elasticity is unlikely to be much above 0.1 and may well be zero. We argue that Hall's estimator for the IES is downward biased because the intra-temporal substitution between nondurable consumption goods and durable consumption goods is ignored and because the changes in real interest rates affect user costs of durable goods. We use a two-step procedure that combines a cointegration approach to preference parameter estimation with Hansen and Singleton's (1982) Generalized Method of Moments approach in order to take these effects into account. In contrast to Hall's result, our estimates for the IES are positive and significantly different from zero even when time aggregation is taken into account.

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I. Introduction

Intertemporal substitution by consumers is a central element of many modern macroeconomic and international models. For example, many of the policy implications of an endogenous growth model studied by Barro (1990) depends on the assumption that the intertemporal elasticity of substitution is positive. In estimating the intertemporal elasticity of substitution (IES), however, Hall (1988) finds that when time aggregation is taken into account, his point estimates are small and not significantly different from zero. His results suggest that intertemporal substitution by consumers is not empirically important.

Hall assumes that preferences are additively separable in nondurable and durable goods, but there is empirical evidence against this assumption (see, e.g., Eichenbaum and Hansen (1990)). When two goods are not additively separable, ignoring one good does not necessarily induce a downward bias in an estimator of the IES for the other good. In the case of nondurable durable goods, however, when the durable good is ignored, the estimators for the IES of the nondurable good are likely to be biased downward. The reason for this is twofold. First, consumption of durable goods is more volatile than nondurable good consumption. In Section III, we will show that the service flow from the durable good purchase is more volatile than nondurable consumption in the U.S. data. Second, real interest rates affect the user cost for the service flow from the durable good. For example, suppose that the real interest rate rises this year. Other things being equal, this results in a higher user cost for the durable good this year and, thus, consumers will substitute away from the durable good and increase today's consumption of the nondurable good. As long as the user cost in the next year does not fall to offset this effect, the growth rate of nondurable consumption decreases compared with the case of no change in user cost. Hence, the estimator of the intertemporal elasticity of substitution which is based only on the growth rate of nondurable consumption growth will be biased downward.

In order to see if this downward bias is important, we use Cooley and Ogaki's (1995) Cointegration-Euler Equation approach, and allow for nonseparable preferences in nondurable and durable goods. We assume that the Constant Elasticity of Substitution (CES) utility function represents intra-temporal preferences.[1] The CES utility function is estimated by a cointegration regression in the first step. In the second step, GMM is applied to the Euler equation with the estimated CES utility function.

Mankiw (1985) estimated the IES for consumption of durable goods. Our approach differs from Mankiw's in that our main focus is on the nonseparability of preferences in nondurable and durable goods while Mankiw assumes separability. However, Mankiw's result that his estimate of the intertemporal elasticity of substitution of durable good consumption is larger than that of nondurable consumption does suggest that the service flow from the durable good purchase is more volatile than nondurable consumption.

Dunn and Singleton (1986), Eichenbaum and Hansen (1990), and Fauvel and Samson (1991) estimate the parameters of Euler equations in models that allow for the nonseparability of preferences between nondurables and durables, though they do not focus on the bias in the estimates of the

[1]Ostry and Reinhart (1992) essentially applied the Cointegration-Euler Equation approach to the CES utility function.

intertemporal elasticity of substitution. The main difference between our paper and theirs is that their estimation method does not allow for adjustment and transactions costs for durable goods, while our method is robust to various forms of these costs. Adjustment and transactions costs are important determinants of of durable good consumption (see, e.g., Bernanke (1985), Lam (1989), and Eberly (1994)). In estimating the intratemporal elasticity in the first step, we use a cointegrating regression which only utilizes long-run information. Hence, as long as adjustment and/or transactions costs do not affect the long-run behavior of durable good consumption, our estimator is consistent. In the second-step GMM estimation, we use the Euler equation obtained by considering changes in nondurable consumption, but not that for changes in durable consumption. It should be noted that the former Euler equation is robust to various forms of adjustment and transactions costs for durable good consumption.

The rest of this paper is organized as follows. Section II presents our theoretical framework for nonseparable preferences in nondurable and durable consumption. Section III explains the data and reports summary statistics, while Section IV explains our econometric method. Section V contains empirical results, and Section VI provides our concluding remarks.

II. Theoretical Framework

In this section, we introduce our model of nonseparable preferences between nondurable and durable consumption. Suppose that a representative consumer maximizes the lifetime utility function

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

in a complete market at time 0, where $E_t(\cdot)$ denotes expectations conditional on the information available at time t. The intra-period utility function is assumed to be of the CES form for the nondurable good (good 1) and the durable good (good 2);

$$u(t) = (a\ C_1(t)^{1-1/\varepsilon} + S_2(t)^{1-1/\varepsilon})^{1/(1-1/\varepsilon)} \qquad (2)$$

where $S_2(t)$ is the service flow from the purchases of good 2. Purchases of the durable consumption good and the service flow are related by

$$S_2(t) = C_2(t) + \delta C_2(t\text{-}1) + \delta^2 C_2(t\text{-}2) + \ldots \qquad (3)$$

where $C_2(t)$ is the real consumption expenditure for good 2 at time t.

Let $P_i(t)$ be the purchase price of consumption good i. We take good 1 as a numeraire for each period: $P_1(t)\equiv 1$. Let $R(t+1)$ be the (gross) return on any asset in terms of good 1, which is realized at $t+1$. Then, the Euler equation is:

$$E_t[\beta R(t+1)mu(t+1)/mu(t)] = 1 \qquad (4),$$

where

$$mu(t) = C_1(t)^{-1/\varepsilon}(a\ C_1(t)^{1-1/\varepsilon} + S_2(t)^{1-1/\varepsilon})^{(\sigma-\varepsilon)/[\sigma(\varepsilon-1)]} \qquad (5).$$

The user cost for the service flow of good 2, $Q(t)$, is:

$$Q(t) = P_2(t) - \delta E_t[\beta P_2(t+1)mu(t+1)/mu(t)]. \qquad (6)$$

Because this formula involves the conditional expectation operator, it is complicated to calculate the user cost. For this reason, we will derive a cointegration restriction which is based on the purchase price, $P_2(t)$ rather than on user cost. We will then use the cointegration restriction to estimate the intraperiod elasticity, ε.

However, it is useful to calculate a proxy for the user cost because one reason for the downward bias in the single good model is that we expect a positive correlation between the user cost and the real interest rate. For the purpose of obtaining a proxy for the user cost, imagine that $mu(t)$ is constant (perhaps because the consumer is risk neutral). Then $E_t[R_{t+1}] = 1/\beta$ from (4), and from (6), the user cost will be

$$Q(t) = P_2(t) - \delta E_t[P_2(t+1)]/E_t[R(t+1)]. \tag{7}$$

We will use (7) to obtain summary statistics of the user cost in the next section.

The following first order condition that states that the user cost is equal to the marginal rate of substitution of the service flow of good 2 and consumption of good 1:

$$Q(t) = a^{-1}[S_2(t)/C_1(t)]^{-1/\varepsilon} \tag{8}$$

In order to derive the restrictions that imply cointegration, it is useful to observe another first order condition which states that the purchase price relative to the price of the nondurable good, $P_2(t)$, is equated with the marginal rate of substitution based on purchases of goods:

$$P_2(t) = \frac{\partial U/\partial C_2(t)}{\partial U/\partial C_1(t)} = \frac{E_t[\Sigma_{\tau=0}^{\infty} \beta^{\tau}\delta^{\tau} mu_2(t+\tau)]}{mu(t)} \tag{9}$$

where

$$mu_2(t) = S_2(t)^{-1/\varepsilon}(a\ C_1(t)^{1-1/\varepsilon} + S_2(t)^{1-1/\varepsilon})^{(\sigma-\varepsilon)/[\sigma(\varepsilon-1)]} \qquad (10).$$

This first order condition forms the basis of the cointegration approach and summarizes the information from the demand side. In order to model the supply side in the simplest way, we consider an endowment economy without production. Let $C_i^*(t)$ be the endowment of good i and $c_i^*(t)=\log(C_i^*(t))$. In an equilibrium, $c_i(t)=\log(C_i(t))=c_i^*(t)$. In a production economy, we require that equilibrium consumption satisfies the assumed trend properties of $c_i^*(t)$. The trend properties of equilibrium consumption are likely to be closely related to those of the technology shock to the good i industry in a production economy.

Assume that $c_i(t)$ is difference stationary for $i=1,2$. Then, the first order condition (9) implies that $P_2(t)\ [C_2(t)/C_1(t)]^{1/\varepsilon}$ is stationary. This follows from the fact that (9) can be used to express $P_2(t)\ [C_2(t)/C_1(t)]^{1/\varepsilon}$ as a function of the stationary variables, $C_1(t+\tau)/C_1(t)$ and $C_2(t+\tau)/C_2(t)$.

III. Data and Summary Statistics

In this section, we explain the data and report summary statistics. We present results for both annual data covering 1929 to 1990 and quarterly data covering 1947:1 to 1990:4. For good 1, we use either nondurables (ND) or nondurables plus services (NDS) from the National Income and Product Account (NIPA). For good 2, we use real durables from the NIPA for the annual data and for the quarterly data either real durables in the NIPA or real durables from Gordon's (1990) data. Gordon's data treats the quality improvement of durable goods in an arguably better way than the NIPA data. We use the implicit deflators as the purchase prices. Because Gordon's data are annual, we use the quarterly series that Ogaki and Park construct from Gordon's data. In constructing the service flow series for durables, (3) is used with the initial condition on $S(t)$ from Musgrave (1979). In Musgrave's data, the depreciation rate is about 18 percent. Wykoff (1970) estimates a depreciation of about 20 percent per year using resale values of automobiles. For our base results, we use $\delta=0.8$ for the annual data and

$\delta=0.94$ for the quarterly data. In order to obtain per capita real consumption, we use resident population for the annual data, and for the quarterly data total population including armed forces overseas (averaged over each quarter).

Nominal interest rate data, together with Barro's average marginal tax rate series, are used to construct nominal after tax rates. These are converted into real rates by the implicit deflator for good 1. For the annual data, we use the six-month commercial paper rate, which is compounded to calculate the one-year rate of return. For the quarterly data, we use the 3-month Treasury Bill rate. Both rates are measured at the end of each period.[2]

Table 1 reports summary statistics for the data. The first panel corresponds to the annual data and the second panel to the quarterly data. In each panel, we first report the standard deviation of the growth rates of consumption and the service flow of durable good purchases. Here the growth rate of each variable is calculated as the first difference of the log of the variable. We note that in both data sets the growth rate of durable consumption much more volatile than that of nondurable consumption. The more relevant comparison for our purpose, however, is between nondurable consumption and the service flow from durable good purchases. The service flow is much smoother than the durable good purchase, but is still much more

[2]We treat the time aggregation problem by lagging the instrumental variables by two periods and allowing the disturbance term to have a one period serial correlation. This does not completely remove the time aggregation problem in our nonlinear model. It should be noted that neither our method nor Hall's (1988) method (which is similar to ours) is perfect. Even in Hall's linear model, we only observe the time average of the level of consumption rather than the time average of the log of consumption. For this reason, we try to avoid further time aggregation problems by using the point-in-time data of the interest rate rather than the time-averaged data.

volatile than nondurable consumption in both data sets. Therefore, ignoring durable good consumption is likely to cause a downward bias in the estimation of the intertemporal elasticity of substitution for total consumption expenditure.

In each panel of Table 1, we also report the standard deviation of the growth rate of the user cost of durable goods relative to a nondurable good price (either the price of ND or of NDS). The standard deviation is positive and statistically significant in all cases. Hence, there may be substantial bias in the estimation of the IES with the single good model when it is applied to total consumption (calculated by adding up nondurable consumption and the service flow from durable good purchases). Hicks's aggregation does not apply when the relative price is not constant.

Table 1 also reports the correlation between the user cost and the real interest rate. We use a Vector Autoregression (VAR) with three lags for the realized real interest rate and the growth rate of the purchase price of the durable good to obtain the expected values of these variables for the calculation of the user cost. We report the correlation of $ln(Q_t)\text{-}ln(Q_{t+1})$ with the expected real interest rate. When this correlation is positive, the estimator of the intertemporal elasticity of substitution is likely to be biased downward as discussed in the Introduction. This correlation is estimated to be positive and significant at the five percent level for both the annual data and quarterly NIPA data. In the case of Gordon's data, the point estimates are positive, but are not statistically significant.

IV. Estimation and Inference

In this section we describe our econometric method. We use Cooley and

Ogaki's (1995) two-step procedure which combines Ogaki and Park's (1989) cointegration approach to preference parameter estimation with Hansen and Singleton's (1982) GMM approach.[3]

A. Implications of the Intratemporal First Order Condition

The notions of stochastic and deterministic cointegration are useful when the economic variables of interest are modeled as difference stationary with drift.[4] This paper focuses on processes that are integrated of order one. Suppose that the components of a vector series $X(t)$ are difference stationary with drift. If a linear combination of $X(t)$, $\gamma' X(t)$ is trend stationary, the components of $X(t)$ are said to be (stochastically) cointegrated, with a cointegrating vector γ. Consider the additional restriction that the cointegrating vector eliminates the deterministic trends as well as the stochastic trends, so that $\gamma' X(t)$ is stationary. This restriction is called the deterministic cointegration restriction.

We assume that the log of equilibrium consumption is difference stationary with drift.[5] Then the cointegration restriction that we derived implies that the log of the relative price and the log of the ratio of nondurable and durable consumption are cointegrated with the deterministic

[3]A similar procedure was used by Ostry and Reinhart (1992) independently.

[4]The notions of stochastic cointegration and the deterministic cointegration restrictions were defined by Ogaki and Park (1989) and Campbell and Perron (1991). Efficiency gains in the estimation of the cointegrating vectors from the imposition of the deterministic cointegration restriction was discussed by West (1989) for the one stochastic trend case and by Hansen (1992) and Park (1992) for the general multiple regressors case.

[5]As shown by Hall (1978), consumption is a random walk when the real interest rate is assumed to be constant. Since we allow the real interest rate to vary over time, the first difference of the log of consumption can have any serial correlation.

cointegration restriction.

B. Step 1: Cointegration

This subsection describes our econometric procedure for the estimation of the cointegrating regression. This procedure allows us to test the null hypothesis of stochastic cointegration and the deterministic cointegration restriction.

Let $X(t)$ be a 2-dimensional difference stationary process: $X(t) - X(t-1) = \phi + \varepsilon(t)$ for $t \geq 1$, where ϕ is a 2-dimensional vector of real numbers, $\varepsilon(t)$ is a stationary process with mean zero, and each component of $\varepsilon(t)$ has a positive long run variance. Suppose that the $X(t)$ are cointegrated, with a cointegrating vector $(1,-\gamma)$, and that the deterministic cointegration restriction is satisfied. Then we can apply Park's (1992) Canonical Cointegrating Regressions (*CCR*) procedure[6] to

$$X_1(t) = \theta_c + \gamma X_2(t) + \varepsilon_c(t). \tag{12}$$

The *CCR* procedure requires us to transform the data before running a regression and corrects for endogeneity and serial correlation. Let $v(t) = (\varepsilon_c(t), \varepsilon_2(t))$ where $\varepsilon_2(t)$ is the second element of $\varepsilon(t)$. Define $\Phi(i) = E(v(t)v(t-i)')$, $\Sigma=\Phi(0)$, $\Gamma = \sum_{i=0}^{\infty}\Phi(i)$, and $\Omega = \sum_{i=-\infty}^{\infty}\Phi(i)$. Here Ω is the long run covariance matrix of v_t. Define

$$\Omega_{11.2} = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21} \tag{13}$$

and $\Gamma_2 = (\Gamma'_{12}, \Gamma'_{22})'$, where Ω_{ij} and Γ_{ij} are the *ij*th component of Ω and Γ, respectively. We make an additional assumption that $\Omega_{11.2}$ is positive. Consider transformations

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[6]See Ogaki (1993a) for a more detailed explanation of CCR based estimation and testing.

$$y^*(t) = y(t) + \Pi_y' v(t) \tag{14}$$

$$X^*(t) = X(t) + \Pi_x' v(t). \tag{15}$$

Because $v(t)$ is stationary, $y^*(t)$ and $X^*(t)$ are cointegrated with the same cointegrating vector $(1, -\gamma)$ as $y(t)$ and $X(t)$ for any Π_y and Π_x. The idea of the CCR is to choose Π_y and Π_x so that the OLS estimator is asymptotically efficient when $y^*(t)$ is regressed on $X^*(t)$. This requires

$$\Pi_y = \Sigma^{-1}\Gamma_2\gamma + (0, \Omega_{12}\Omega_{22}^{-1}) \tag{16}$$

$$\Pi_x = \Sigma^{-1}\Gamma_2. \tag{17}$$

In practice, long-run covariance parameters in these formulas are estimated, and the estimated Π_y and Π_x are used to transform $y(t)$ and $X(t)$. As long as these parameters are estimated consistently, the resultant CCR estimator is asymptotically efficient.

The *CCR* estimators have asymptotic distributions that can essentially be considered normal, implying that their standard errors have the usual interpretation.[7] An important property of the *CCR* procedure is that linear restrictions can be tested by χ^2 tests, which are free from nuisance parameters. We use χ^2 tests in a regression with spurious deterministic

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[7]The CCR estimators are asymptotically efficient, but there are other asymptotically efficient estimators such as those developed by Saikkonen (1989), Phillips and Hansen (1990), Phillips (1991), and Stock and Watson (1993). Johansen's estimators are often used, but Johansen assumes a Gaussian VAR structure. The CCR does not require this Gaussian VAR assumption, which is important for our purpose because our economic model implies nonlinear short-run dynamics. Monte Carlo experiments in Park and Ogaki (1991) show that the CCR estimators have better small sample properties in terms of the mean square error than Johansen's estimators. Following Monte Carlo based recommendations by Park and Ogaki (1991) and Han and Ogaki (1991), we used the prewhitening method and report third stage CCR estimates and fourth stage CCR $H(p,q)$ test statistics.

trends added to (12) in order to test for stochastic and deterministic cointegration. For this purpose, the *CCR* procedure is applied to the regression

$$X_1(t) = \theta_c + \sum_{i=1}^{q} \eta_i t^i + \gamma X_2(t) + \varepsilon_c(t). \tag{18}$$

Let $H(p,q)$ denote the standard Wald statistic under the hypothesis $\eta_p = \eta_{p+1} = \ldots = \eta_q = 0$ with the estimate of the variance of $\varepsilon_c(t)$ replaced by $\Omega_{11.2}$ (see Park (1990) for details). Then $H(p,q)$ converges in distribution to a $\chi^2_{p\text{-}q}$ random variable under the null of cointegration. In particular, the $H(0,1)$ statistic tests the deterministic cointegrating restriction. On the other hand, the $H(1,q)$ statistic tests stochastic cointegration.

C. Step 2: The Estimation of the Intertemporal Elasticity of Substitution

In Step 1, we obtain a consistent estimate of the intratemporal elasticity, ε. The second step of our procedure is to apply GMM to the Euler equation (4) in order to obtain estimates of intertemporal parameters. This two-step procedure does not alter the asymptotic distributions of the GMM estimators and test statistics because our cointegrating regression estimator is super consistent and converges at a rate faster than $T^{1/2}$.

The time aggregation problem is handled by lagging the instrumental variables two periods and by allowing the disturbance to have an a moving average of order one (MA(1)) structure in the calculation of the optimum weighting matrix. Hall's econometric method assumes that the MA coefficient for the disturbance is known, but we estimate the MA coefficient in the GMM framework. We do not make the assumption that the MA coefficient is known because the value of the coefficient can deviate from the value that Hall's theory predicts: for example, the planning period of the consumer may be different from the one assumed by Hall.

V. Empirical Results

This section reports the results of the cointegrating regressions from Step 1 for intratemporal first order condition and the Step 2 GMM estimation of the Euler equation.

A. Annual Data

Table 2 reports the cointegrating regression results based on CCR for ND and NDS with and without the dummy variable for 1940-45 for World War II (WWII).[8] For ND, the dummy variable is significant at the five percent level. For NDS, the dummy variable is not significant at the five percent level, but is significant at the ten percent level. In addition, the H(p,q) tests are more favorable for the specification with the dummy variable. Among the four H(p,q) test statistics reported for ND with the dummy variable, only one is significant at the ten percent level and none of them is significant at the one percent level. Among the four H(p,q) test statistics for NDS with the dummy variable, one is marginally significant at the one percent level and another is significant at the five percent level. Overall, the evidence against cointegration is not strong because the H(p,q) tests often overreject according to Han and Ogaki (1991).

For all cases, the intratemporal elasticity of substitution, ε, is

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[8]We used Ogaki's (1993c) GAUSS CCR Package for the CCR estimation. The CCR procedure requires an estimate of the long run covariance of the disturbances in the system. We used Park and Ogaki's (1991) method with Andrews and Monahan's (1992) prewhitened HAC estimator with the QS kernel. A VAR of order one was used for prewhitening. We followed footnote 4 of Andrews and Monahan and the maximum absolute value of the elements of Δ notation was set to 0.99. Andrews's (1991) automatic bandwidth estimator, S_T, was constructed from fitting AR(1) to each disturbance.

estimated with the theoretically correct positive sign. For ND, the intertemporal elasticity of substitution is also estimated to be significantly larger than one at the five percent level, so that the Cobb-Douglas utility function is rejected. For NDS, our point estimates for ε are not significantly different from either zero or one.

Table 3 presents the GMM results.[9] The instrumental variables are a constant, the realized real interest rate, the growth rate of the real consumption ratio of good 1 and good 2, and the real defense expenditure growth rate. All instruments are lagged two periods rather than one. Including the growth rate of consumption of good 1, which is often used as an instrument, led to convergence problems after one or two iterations. This fact and Hall's (1978) finding that consumption growth has, at most, only weak serial correlation suggest that the growth rate of consumption of good 1 is not a good instrument.

The first panel presents our results for the two-good model described in Section II. The second panel presents our results for the one-good model, which can be obtained by assuming $\sigma=\varepsilon$ (which is the separability case). For the one-good model, *a* is normalized to one. While the one-good model is similar to Hall's (1988) model, we include the results because the econometric method and sample period are somewhat different. Unlike Hall, we do not linearize the Euler equation (4) due to the difficulty in doing so for the two-good model. We use exactly the same econometric method and data

[9]We used Hansen-Heaton-Ogaki GAUSS GMM package described in Ogaki (1993bd) that was supported by NSF Grants SES-3512371 and SES-9213930 for the GMM estimation. We iterated on the weighting matrix as described by Kocherlakota (1990) up to four iterations, since his Monte Carlo results indicated that the iteration improves the small sample properties of the GMM estimator.

for both the one-good and two-good models, so that we can directly compare the results.

In all cases, Hansen's J test of the overidentifying restrictions does not reject the model at the conventional levels. For both ND and NDS, our point estimates of σ are positive and significantly different from zero at the five percent level for the two-good model. In contrast, the one-good model yields smaller point estimates of σ for both ND and NDS with similar standard errors. It should be noted that the separability assumption ($\sigma=\varepsilon$) is rejected in the two-good model for both ND and NDS.

B. Quarterly Data

Table 4 reports the cointegrating regression results from Step 1 based on CCR. For ND and NIPA durables goods, when the full sample period of 1947:2-1990:4 is used, the H(0,1) test implies a very strong rejection of the model. This is likely to be due to the fact that the level of the stock of durable goods was very low immediately after WWII. This meant that the stock of durable goods grew faster than the purchases of durable goods in the period immediately following WWII. Because the cointegration regression depends on the assumption that the stock of durable goods grows at the same rate as the purchase of durable goods, it is appropriate to start the sample period at a later date. For the sample period of 1951:1-1990:4, none of the H(p,q) test statistics is large enough to reject the model at the conventional levels.

For ND and Gordon's durables data, the H(0,1) test rejects the model at the five percent level when the full sample period of 1947:2-1983:4 is used. Again, none of the H(p,q) test statistics rejects the model at the conventional levels when the sample period of 1951:1-1983:4 is used.

In all cases, the intratemporal elasticity of substitution, ε, is estimated with the theoretically correct positive sign and is significantly different from zero. The intertemporal elasticity of substitution, σ, is also estimated to be significantly greater than one at the five percent level, so that the Cobb-Douglas utility function is rejected.

For NDS and NIPA durables, none of the H(p,q) test statistics is significant at the twenty percent level over either sample period. In addition, the point estimates of ε are similar for both sample periods. The estimates of ε have the theoretically correct positive sign and are significantly different from zero. The intertemporal elasticity of substitution is estimated to be significantly smaller than one, so that the Cobb-Douglas utility function is rejected.

For NDS and Gordon's durables, none of the H(p,q) test statistics reject the model at the conventional levels. In contrast to the previous results, the Cobb-Douglas utility function is not rejected at the five percent level.

Table 5 presents the GMM results. The instrumental variables are the same as those used for the annual data plus a yield spread. The yield spread is the monthly yield to maturity of corporate bonds rated Baa by Moody's Investor Services, minus the Aaa corporate yield. All instruments are lagged two periods rather than one. We were not able to obtain convergence when β is estimated with σ. Therefore, we report results when β is fixed.[10]

[10]For the estimation of the results of Table 5, we penalize the exceptionally high values of σ. For this purpose, we multiply the

The first panel presents our results for the two-good model, and the second panel presents those for the one-good model. In the case of two goods, Hansen's J test rejects the model at the one percent level both for ND and NDS when the NIPA durable good data are used and $\beta=0.990$. On the other hand, the J-test does not reject the model at the five percent level for ND, and NDS when Gordon's data are used and $\beta=0.990$. We also report results for $\beta=0.995$ and $\beta=0.985$ for Gordon's data. In an economy without growth, $\beta=0.990$ implies a real interest rate of about 4.1 percent; $\beta=0.995$, about 2 percent; and $\beta=0.985$, about 6 percent.

In all cases of the two-good model, our point estimates of σ are positive and significantly different from zero at the five percent level. In contrast, the one-good model yields negative point estimates of σ for both ND and NDS. The separability assumption ($\sigma=\varepsilon$) is rejected in the two-good model for ND except for the case of $\beta=0.985$. For NDS, the separability assumption is rejected at the ten percent level with $\beta=0.990$ and at the five percent level with $\beta=0.995$. However, the assumption cannot be rejected with $\beta=0.985$. Thus, the evidence against separability is mixed for the quarterly data.

VI. Conclusions

In this paper, we have argued that ignoring the intratemporal substitution between nondurables and durables is likely to lead to a downward bias in an estimate of the IES. When we account for this

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disturbance term $(|\sigma|-10)^2$ when the absolute value of σ used in the nonlinear search program is greater than ten. This bound was sometimes reached in earlier iterations for the weighting matrix but was never reached in the last iteration.

intratemporal substitution, our empirical results are very different from those of Hall, who concludes that "the elasticity is unlikely to be much above 0.1 and may well be zero". The IES is estimated to be positive and significant, and the point estimates of the IES under the nonseparability assumption range from 0.414 to 1.156. In contrast, the point estimates of the IES based on our one-good model under the separability assumption are sometimes negative and are always smaller than the corresponding point estimates under the nonseparability assumption.

We found strong empirical evidence against separability of preferences between ND and durable goods. We found evidence against separability between NDS and durable goods in the annual data, though evidence for the quarterly data was mixed. In particular, our empirical results indicate that the intratemporal elasticity between ND and durable goods is much higher than the intertemporal elasticity of substitution. This finding, together with the fact that part of durable good purchases are theoretically one type of saving, suggests that some of the puzzling behavior we observe with regard to saving may be explained by the addition of the intratemporal substitution between nondurable and durable consumption goods to standard models of saving.

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Table 1

SUMMARY STATISTICS

Annual Data	Statistics	S.E.
Standard Deviation of Growth Rates		
Durable Good Consumption	0.1315	0.0248
Durable Good Service Flow	0.0394	0.0077
Nondurable Good Consumption (ND)	0.0279	0.0063
Nondurable Good Consumption (NDS)	0.0249	0.0051
Standard Deviation of the User Cost		
ND as numeraire	0.2231	0.0841
NDS as numeraire	0.1644	0.0564
Correlation: User Cost and Real Interest Rate		
ND as numeraire	0.4664	0.0673
NDS as numeraire	0.4528	0.1380

Quarterly Data		Statistics	S.E.
Standard Deviation of Growth Rates			
	Durable Good Consumption (NIPA)	0.0423	0.0045
	Durable Good Service Flow (NIPA)	0.0276	0.0016
	Durable Good Consumption (Gordon)	0.0523	0.0087
	Durable Good Service Flow (Gordon)	0.0176	0.0026
	Nondurable Good Consumption (ND)	0.0085	0.0006
	Nondurable Good Consumption (NDS)	0.0059	0.0005
Standard Deviation of the User Cost			
	NIPA Data, ND as numeraire	0.0644	0.0127
	NIPA Data, NDS as numeraire	0.0595	0.0093
	Gordon's Data, ND as numeraire	0.0817	0.0149
	Gordon's Data, NDS as numeraire	0.0178	0.0066
Correlation: User Cost and Real Interest Rate			
	NIPA Data, ND as numeraire	0.1561	0.0573
	NIPA Data, NDS as numeraire	0.2010	0.0717
	Gordon's Data, ND as numeraire	0.1143	0.1031
	Gordon's Data, NDS as numeraire	0.1574	0.0988

NOTE: The standard errors are calculated by COREST.EXP program in Hansen/Heaton/Ogaki's GAUSS GMM package, using a VAR(1) prewhitened QS kernel estimator with Andrews's (1991) automatic bandwidth selection. We use $\delta=0.94$ to calculate the service flows and the user cost of durable goods.

TABLE 2

CANONICAL COINTEGRATING REGRESSION RESULTS
FOR ANNUAL DATA

Nondurable Good	ε	d	$H(0,1)$	$H(1,2)$	$H(1,3)$	$H(1,4)$
(1)	(2)	(3)	(4)	(5)	(6)	(7)
ND	3.951	. . .	1.531	0.208	0.298	2.719
	(1.329)		(0.216)	(0.648)	(0.862)	(0.437)
ND	2.861	0.711	5.918	1.119	1.128	2.518
	(0.807)	(0.269)	(0.015)	(0.290)	(0.569)	(0.472)
NDS	0.964	. . .	3.205	6.271	6.366	8.232
	(0.628)		(0.073)	(0.012)	(0.041)	(0.041)
NDS	0.980	0.404	1.407	6.641	6.664	6.924
	(0.535)	(0.225)	(0.236)	(0.010)	(0.036)	(0.074)

NOTE: Park and Ogaki's (1991) method with Andrews's (1991) automatic bandwidth parameter estimator was used to estimate long-run correlation parameters. In cols. 2 and 3, standard errors are in parentheses. Col. 3 gives a coefficient of the dummy variable for WWII when it is included in the regression. Col. 4 is a χ^2 test statistic for the deterministic cointegration restriction. Asymptotic P-values are in parentheses. Cols. 5, 6 and 7 are χ^2 test statistics for stochastic cointegration. Asymptotic P-values are in parentheses.

TABLE 3

GENERALIZED METHOD OF MOMENTS RESULTS
FOR ANNUAL DATA

Nondurable Good Data (1)	ε (2)	σ (3)	β (4)	J_T (5)
The Two-Good Model				
ND	2.861	0.766 (0.340)	1.032 (0.016)	1.778 (0.411)
NDS	0.980	0.414 (0.186)	1.065 (0.032)	1.493 (0.474)
The One-Good Model				
ND		0.588 (0.381)	0.979 (0.019)	2.378 (0.304)
NDS		0.270 (0.183)	0.934 (0.050)	3.319 (0.190)

NOTE: In cols. 3 and 4, standard errors are in parentheses. Col. 5 reports Hansen's J test with two degrees of freedom, with asymptotic P-values in parentheses.

TABLE 4

CANONICAL COINTEGRATING REGRESSION RESULTS FOR QUARTERLY DATA

Nondurable Good	Durable Good	Sample Period	ε	*H(0,1)*	*H(1,2)*	*H(1,3)*	*H(1,4)*
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ND	NIPA	47:2-90:4	1.864 (0.262)	14.125 (0.000)	0.760 (0.383)	0.763 (0.683)	0.763 (0.858)
ND	NIPA	51:1-90:4	1.527 (0.235)	2.583 (0.108)	0.008 (0.928)	0.008 (0.996)	0.008 (0.999)
ND	Gordon	47:2-83:4	1.323 (0.098)	4.958 (0.026)	0.700 (0.403)	1.345 (0.510)	2.400 (0.494)
ND	Gordon	51:1-83:4	1.246 (0.095)	0.455 (0.500)	0.671 (0.413)	1.701 (0.427)	2.400 (0.494)
NDS	NIPA	47:2-90:4	0.794 (0.116)	0.005 (0.945)	0.675 (0.411)	0.707 (0.702)	1.364 (0.714)
NDS	NIPA	51:1-90:4	0.747 (0.114)	0.040 (0.842)	1.566 (0.211)	1.617 (0.446)	4.664 (0.198)
NDS	Gordon	47:2-83:4	0.913 (0.079)	0.947 (0.331)	0.594 (0.441)	1.478 (0.478)	2.742 (0.433)
NDS	Gordon	51:1-83:4	0.879 (0.083)	0.353 (0.552)	0.571 (0.450)	4.356 (0.113)	4.380 (0.223)

NOTE: Park and Ogaki's (1991) method with Andrews's (1991) automatic bandwidth parameter estimator was used to estimate long-run correlation parameters. In col. 4, standard errors are in parentheses. Col. 5 is a χ^2 test statistic for the deterministic cointegration restriction. Asymptotic P-values are in parentheses. Cols. 6, 7 and 8 are χ^2 test statistics for stochastic cointegration. Asymptotic P-values are in parentheses.

TABLE 5

GENERALIZED METHOD OF MOMENTS RESULTS
FOR QUARTERLY DATA

Nondurable Good Data (1)	Durable Good Data (2)	Sample Period	ε (3)	β (4)	σ (5)	J_T (6)
The Two-Good Model						
ND	NIPA	47:2-90:4	1.527	0.990	0.505 (0.094)	13.270 (0.010)
ND	Gordon	47:2-83:4	1.246	0.990	0.662 (0.098)	8.092 (0.088)
ND	Gordon	47:2-83:4	1.246	0.995	0.496 (0.065)	8.519 (0.074)
ND	Gordon	47:2-83:4	1.246	0.985	1.004 (0.187)	7.620 (0.107)
NDS	NIPA	47:2-90:4	0.747	0.990	0.588 (0.089)	15.220 (0.004)
NDS	Gordon	47:2-83:4	0.879	0.990	0.727 (0.106)	9.392 (0.052)
NDS	Gordon	47:2-83:4	0.879	0.995	0.529 (0.065)	9.787 (0.044)
NDS	Gordon	47:2-83:4	0.879	0.985	1.156 (0.231)	8.826 (0.066)
The One-Good Model						
ND	. . .	47:2-90:4	. . .	0.990	-0.214 (0.060)	7.534 (0.057)
ND	. . .	47:2-83:4	. . .	0.990	-0.315 (0.106)	6.063 (0.109)
NDS	. . .	47:2-90:4	. . .	0.990	-0.393 (0.079)	9.719 (0.021)
NDS	. . .	47:2-83:4	. . .	0.990	-0.559 (0.130)	6.949 (0.074)

NOTE: The results in this table are obtained by setting ε to be 1.527 for ND and 0.747 for NDS. In cols. 3 and 4, standard errors are in parentheses. Col. 5 reports Hansen's J test with 8 degrees of freedom, and asymptotic P-values in parentheses.