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Abstract

In this paper, we use long-run annual data to estimate the intertemporal elasticity of substitution while accounting for the intra-temporal substitution between nondurable consumption goods and durable consumption goods. We apply a two-step procedure that combines a cointegration approach to preference parameter estimation with Generalized Method of Moments.

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1. Introduction

As Hall (1988) points out, intertemporal substitution by consumers is a central element of many modem macroeconomic and international models. The quantitative importance of effects of changes in various policies implied by these models depend on the magnitude of the intertemporal elasticity of substitution (IES). However, when time aggregation is taken into account, point estimates of the IES tend to be small or even negative, Hall (1988) finds that when time aggregation is taken into account, his point estimates are small and not significantly different from zero. He concludes that the elasticity is unlikely to be much above 0.1 and may well be zero. Thus, his results suggest that intertemporal substitution by consumers is not empirically important.

Working with a similar economic model, Hansen and Singleton (1996) improved on Hall's inference methods with a technique that is scale invariant and asymptotically efficient. While they find that there is considerably less precision in the estimation and evidence against small positive values of the IES, their point estimates are negative.
In the companion paper, Ogaki and Reinhart (1998), we argue that the model used by these authors is misspecified because the intra-temporal substitution between nondurable consumption goods and durable consumption goods is ignored. Both Hall (1988) and Hansen and Singleton (1996) assume 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)). In principle, when two goods are not additively separable, ignoring one good in estimating the IES of the other good does not necessarily induce a bias that increases the probability of finding either small and positive point estimates or estimates with the wrong sign. In the case of nondurable and durable

goods, however, ignoring durable goods in estimating the IES, as in Hall (1988) and Hansen and Singleton (1996), likely introduces a bias in this direction.

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 additive1y 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 3, 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 (1996) 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. The CES utility function is estimated by Ogaki and Park (1998) co integration approach to estimating preference parameters in the first step. In the second step, Generalized Method of Moments (GMM) is applied to the Euler equation with the estimated CES utility function.

In Ogaki and Reinhart (1998), we apply this two step approach to post war U.S. quarterly data and find that our estimates for the IES are positive and significantly different from zero, even when time aggregation is taken into account. In this paper, we apply the approach to long-run U.S. annual data. It is important to confirm our findings from post war data with long run data because the long-run data are more appropriate for the cointegration approach.

2. 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

(1)

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

in a complete market at time 0, where E_t (.) 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);

(2)

$$u(t) = (aC_1(t)^{1-1/\varepsilon} + S_2(t)^{1-1/\varepsilon})^{1/(1-1/\varepsilon)}$$

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

(3)

$$S_2(t) = C_2(t) + \delta C_2(t-1) + \delta^2 C_2(t-2) + ..$$

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: $P_i\,(t) = 1$. 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:

(4)

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

where

(5)

$$mu(t) = C_1(t)^{-1/\varepsilon}(aC_1(t)^{1-1/\varepsilon} + S_2(t)^{1-1/\varepsilon})^{(\sigma-\varepsilon)/[\sigma(\varepsilon-1)]}$$

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:

(6)

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

where

(7)

$$mu_2(t) = S_2(t)^{-1/\varepsilon}(aC_1(t)^{1-1/\varepsilon} + S_2(t)^{1-1/\varepsilon})^{(\sigma-\varepsilon)/[\sigma(\varepsilon-1)]}$$

This first order condition forms the basis of the cointegration approach and summarizes the information from the demand side: Ogaki and Reinhart (1998) show that under certain conditions, the first order condition (7) implies that $P_2(t)[C_2(t)/C_i\ (t)]$ 1/ ε is stationary.

3. Empirical results

This section explains the data and reports the empirical results of the two step approach. The data are annual and cover 1929 to 1990. 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.

We use the implicit deflators as the purchase prices. In constructing the service flow series for durables, (3) is used with the initial condition on Set) 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 /5=0.8 for the annual data and (5 = 0.94 for the quarterly data. In order to obtain per capita real consumption, we use resident population. 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. We use the six-month commercial paper rate, which is compounded to calculate the one-year rate of return.

In Step I, we apply a cointegrating regression to the intratemporal first order condition (9) in order to estimate the intratemporal elasticity of substitution, B. In Step 2, GMM is applied to the Euler Eq. (4). For details of the econometric method, see Ogaki and Reinhart (1998). Table 1 reports the cointegrating regression results based on Park (1992) Canonical Cointegrating Regression (CCR) for ND and NDS with and without the dummy variable for 1940-45 for World War II (WWII). 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 (1997).

For all cases, the intratemporal elasticity of substitution, B, is estimated with the theoretically correct positive sign. For ND, the intratemporal 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 B are not significantly different from either zero or one.

Table 2 presents the GMM results. 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 (1988) finding that consumption growth has, at most,

Table 1. Canonical cointegrating regression results

Nondurable Good (1)	σ (2)	d (3)	$H(0, 1)$ (4)	$H(1, 2)$ (5)	$H(1, 3)$ (6)	$H(1, 4)$ (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: In columns, 2 and 3, standard errors are in parentheses, column 3 gives a coefficient of the dummy variable for the WWII when it is included in the regression. Column 4 is a X_2 test statistic for the deterministic cointegration restriction, Asymptotic P-values are in parentheses, Cols, 5, 6 and 7 are X_2 test statistics for stochastic cointegration. Asymptotic P-values are in parentheses.

Table 2. Generalized method of moments results

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

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 2. The second panel presents our results for the one-good model, which can be obtained by assuming (T = E: which is the separability case). For the one-good model, a is normalized to one. While the one-good model is similar to Hall (1988) model, we include the results because the econometric method and sample period are somewhat different. Unlike Hall, we do not linearize the Euler Eq. (4) due to the difficulty in doing so for the two-good model. We use exactly the same econometric method and data 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 (T 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 (T for both ND and NDS with similar standard errors. It should be noted that the separability assumption is rejected in the two-good model for both ND and NDS.

4. Conclusions

In this paper, we have used long-run annual data to confirm Ogaki and Reinhart (1998) findings from postwar quarterly data. This task is important because we estimate the intratemporal elasticity of substitution from a cointegrating regression in our two step procedure. Because cointegration is a long-run relationship, it is desirable to use long-run data. Our results from the long-run data are similar to those from postwar data. The IES is estimated to be positive and significant when the role of durable good consumption is taken into account.

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