Estimating the substitutability between private and public consumption: the case of Spain, 1960-2001[*]

Vicente Esteve[†]

Department of Applied Economics II,
University of Valencia, 46071, Valencia, Spain

Juan Sanchis

Department of Applied Economics II,
University of Valencia, 46071, Valencia, Spain

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Abstract

This paper examines the relationship between private and public consumption using Spanish data over the period 1960-2001. We extend the previous analysis addressing the question of whether this relationship is stable over time, or exhibit a structural break allowing the instability to occur at an unknown date. Our empirical results are consistent with the existence of a long-run relationship between private and public consumption. Thus, we have been detected the structural changes or regime shifts in the cointegration regression around the time of the oil price shock of 1973/74. In addition, the estimates both the intratemporal and intertemporal elasticities of substitution between the two types of expenditure (in full-sample and in both sub-samples) suggests that private and public consumption in Spanish economy are Edgeworth-Pareto substitutes.

Keywords: Fiscal policy, government expenditure, private consumption, structural breaks, cointegration, GMM.

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[†]Corresponding author: Vicente Esteve, Departamento de Economía Aplicada II, Universidad de Valencia, Apartado Oficial 22006, 46071 Valencia, Spain. E-mail: vicente.esteve@uv.es.

1 Introduction

The fcreation of the Economic and Monetary Union (EMU) by 12 member countries of the European Union (EU) implied that …scal policy became the main instrument of stabilization policies available to their national authorities. As a consequence, issues related with …scal policy have gained a growing interest in last years, in both academic and policy circles. In particular, the long-run relationship between private and government consumption has become a matter of great concern for its implications for the process of …scal consolidation during the 1990s. This consolidation has been planned in the majority of EU countries to fall almost entirely on the expenditure side of the budget. Whether this process of budget cuts will have a short-run impact on real economic activity depends basically on the private sector's willingness to substitute its own expenditure for public consumption. For example, if this substitution e¤ect exists, it will dampen the e¤ect on demand of cut-backs to government consumption. The possible relationship of substitutability between public and private consumption has important implications for assessing the overall e¤ectiveness of …scal policy. If substitutability is empirically accepted, it would support the theoretical insights put forward by Giavazzi and Pagano (1990, 1996) and other authors, who introduced the possibility of "expansionary e¤ects of …scal contractions" or the "non-keynesian e¤ects" of the …scal policy.[1]

In the empirical application we use data for Spain, a country traditionally experiencing high budget de…cits, which has accomplished an important …scal consolidation in last years. Such e¤orts were necessary to satisfy the requirements set in the Treaty of Maastricht, in order to be able to participate in the Economic and Monetary Union (EMU) launched in Europe after 1999.

The impact of the government consumption depend upon whether this variable increase or decrease the marginal utility of private consumption, i.e. whether government consumption are Edgeworth-Pareto complements or substitutes for private consumption. Unfortunately, the empirical works on the substitutability between private and public consumption are still inconclusive due to di¤erences in the econometric methodology, the particular speci…cation of the private agent's utility function, and the sample period used.[2]

In this paper we examine the relationship between private and public consumption using Spanish data over the period 1960-2001. Previous research for Spanish economy has found either that private and public con-

[1]For a survey of the abundant theoretical and empirical literature on non-keynesian e¤ects of the …scal policy, and over the European experience of …scal consolidations see Zaghini (2001).

[2]For a recent survey of the abundant theoretical and empirical literature on degree of substitutability between government spending and private consumption see Ho (2001).

sumption are substitutes (Esteve et al. (1997)) or that private sector consumption is unaffected by government consumption (Marchante (1993)). We update these works by incorporating a new model developed by Amano and Wirjanto (1998) which allow us to estimate both the intratemporal and intertemporal elasticities of substitution for private and public consumption and by adopting and econometric approach which allow us to detect whether the relationship between private and public consumption is stable over time, or exhibit a structural break allowing the instability to occur at an unknown date.

Our empirical results are consistent with the existence of a long-run relationship between private and public consumption in Spanish economy. Thus, we have been detected the structural changes or regime shifts in the cointegration regression around the time of the oil price shock of 1973/74. In addition, the estimates both the intratemporal and intertemporal elasticities of substitution between the two types of expenditure (in full-sample and in both sub-samples) suggests that private and public consumption in Spanish economy are Edgeworth-Pareto substitutes, suggesting that an decrease in public consumption will induce an increase in consumption by the private sector. These results support the theoretical insights put forward by Giavazzi and Pagano (1990, 1996) and other authors, who introduced the possibility of "expansionary effects of fiscal contractions" or the "non-keynesian effects" of the fiscal policy.

The rest of the paper is organized as follows: The theoretical framework is presented in Section 2. Section 3 describes the data, and present the estimation results from Spanish economy. Finally, Section 4 summarizes draws the conclusions.

2 The theoretical model and its testable implications

We use a two-good permanent-income model in which the expected lifetime utility for a representative consumer is assumed to be

$$E_0 \left[\sum_{t=0}^{\infty} \beta^t U(C_t; G_t) \right] \tag{1}$$

where E_t is the conditional expectations operator in period t, β is an intertemporal discount factor, C_t is real private consumption, G_t is real public consumption, and $U(.)$ is a concave intraperiod utility function for private and public consumption.[3]

Following Amano and Wirjanto (1998) we consider a constant-elasticity-of-substitution utility function

[3] A similar approach is taken in Amano and Wirjanto (1998).

$$U(C_t;G_t) = \frac{u(C_t;G_t)^{1-\gamma}}{1-\gamma}; \quad 1/\gamma > 0 \text{ and } 1/\gamma \neq 1 \tag{2}$$

where $1/\gamma$ represents the intertemporal elasticity of substitution for consumption. For $1/\gamma = 1$ we define $U(C_t;G_t)$ in logarithms.

We assume that the intraperiod utility function takes the following form

$$u(C_t;G_t) = [\phi C_t^{1-\alpha} + (1-\phi)G_t^{1-\alpha}]^{1/(1-\alpha)}; \quad 1/\alpha > 0 \text{ and } 1/\alpha \neq 1 \tag{3}$$

where $1/\alpha$ represents the intratemporal elasticity of substitution between private and public consumption, and ϕ and $(1-\phi)$ are the weights of private and public consumption, respectively.[4]

Using the above specifications Amano and Wirjanto (1998) proposed three testable implications based on the fact that the substitutability between C_t and G_t (i.e.$U_{CG;t}$), depends on the relative magnitude of the intertemporal and intratemporal elasticities of substitution.[5] These implications are the following:[6]

(1) If $1/\gamma > 1/\alpha$ then C_t and G_t are Edgeworth-Pareto complements.

(2) If $1/\gamma < 1/\alpha$ then C_t and G_t are Edgeworth-Pareto substitutes.

(3) and if $1/\gamma = 1/\alpha$ then they are Edgeworth independent or unrelated.

Lets define P_t as the relative price of public consumption with respect to private consumption. One of the intraperiod first-order conditions of the model equates this relative price to the marginal rate of substitution on the purchase of the two goods,

$$P_t = (\partial U/\partial G_t)/(\partial U/\partial C_t) \tag{4}$$

Assuming that the logarithm of the consumption series is an I(1) process with drift and P_t is measured with stationary measurement error, then a cointegrating regression implied by the equation (4) is given by

$$\ln P_t = \mu + \alpha \ln(C_t/G_t) + \varepsilon_t \tag{5}$$

where $\mu = \ln[(1-\phi)/\phi]$, the error term is an I(0) process with mean zero and $(1; -\alpha)$ is the cointegrating vector. The appropriate estimation of equation (5) would allow to obtain a consistent estimate of the intratemporal elasticity of substitution between private and public consumption, $1/\alpha$. Furthermore, we would estimate the intertemporal elasticity of substitution, $1/\gamma$, and the discount factor, β, in a second stage, by impossing the above estimate in the Euler equation.

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[4] For $1/\alpha = 1$ we assume that $u(C_t;G_t) = C_t^{\phi} G_t^{1-\phi}$.

[5] Amano and Wirjanto (1998) show that the sign of the cross second partial derivates of $U_{CG;t}$ is determined by the relative magnitute of both elasticities. Specifically, the $\text{sign}[U_{CG;t}] = \text{sign}[1/\gamma - 1/\alpha]$.

[6] For more details, see Newman (1998).

Let R_{t+1} be the gross return on any asset between t and $t+1$ (or the real interest rate) expressed in units of the private good. Then the intertemporal first-order condition or Euler equation for private consumption is

$$E_t \beta [(\partial U/\partial C_{t+1})/(\partial U/\partial C_t)] R_{t+1} = 1 \tag{6}$$

Equation (6) is estimated using the generalized method of moments (GMM) procedure proposed by Hansen (1982) and Hansen and Singleton (1982), defining the estimation funciton as

$$\xi_{t+1} = \beta [(\partial U/\partial C_{t+1})/(\partial U/\partial C_t)] R_{t+1} - 1 \tag{7}$$

and it is assumed that $E_t[\xi_{t+1}] = 0$.[7]

3 Empirical results

The estimation of the relationship between private and public consumption is undertaken in two steps. In the first step, we estimate the intratemporal elasticity of substitution, $1/\alpha$, using a cointegration approach, which allows us to avoid any spurious regression and to retain the long-run information. To do so, we first test for unit roots to determine the order of integration of the two series (equation (5); before the estimatation of the cointegration relationship between the variables. Once we estimate the intratemporal elasticity of substitution, we analyze the possibility of structural changes in the above relationship. In the second step, we estimate the intertemporal elasticity of substitution, $1/\gamma$, and the discount factor, β, using the GMM procedure.

3.1 Data

We use Spanish annual data for the period 1960-2001 drawn from the Bank of Spain. The variable of interes, the rate C_t/G_t,is obtained using publice and privat final consumption expenditure at 1995 prices. The implicit prices for public and private expenditure are constructed by dividing nominal expenditures by their 1995 constant euro counterpart. The relative price, P_t, is calculated as the ratio of these two price indexes. We approximate the gross return by the real long term interest rate, $R_t = i_t - \pi_t$, where i_t is the nominal long-term interest rate (private bonds of electric utilities before February 1978; from March 1978 to December 1992, central government bonds at more than two years; and, from January 1993, central government benchmark bond of 10 years), and π_t is the annual change of the private consumption implicit price index (1995=100). In the empirical application lower case letters indicate the natural log of a variable.

[7] We thus exploit this this moment condition to implement the GMM procedure, i.e. $E[Z_t v_{t+1}] = 0$ where Z_t is a set of instrumental variables.

3.2 Unit root tests

As stated above, we ...rst examine the properties of the series. To do so, ee use a modi...ed version of the Dickey and Fuller (1979, 1981) test (DF) and a modi...ed version of the Phillips and Perron (1988) tests (PP) proposed by Ng and Perron (2001) for the null of a unit root.

In general, the majority of the conventional unit root tests (DF and PP types) su¤er from three problems. First, many tests have low power when the root of the autoregressive polynomial is close to but less than unit, Dejong et al. (1992). Second, the majority of the tests su¤er from severe size distortions when the moving-average polynomial of the ...rst di¤erences series has a large negative autoregressive root, Schwert (1989) and Perron and Ng (1996). Third, the implementation of unit root tests often necessitates the selection of an autoregressive truncation lag, k. However, as discussed in Ng and Perron (1995) there is a strong association between k and the severity of size distortions and/or the extend of power loss. Recently, Ng and Perron (2001) proposed a methodology that solves these three problems. This method consists of a class of modi...ed tests, called $\bar{M}^{GLS}_{MAIC}$, originally developed in Stock (1999) as M tests, with GLS detrending of the data as proposed in Elliot et al. (1996), and using the Modi...ed Akaike Information Criteria (MAIC).[8] Also, Ng and Perron (2001) have proposed a similar procedure to correct for the problems of the standard Augmented Dickey-Fuller (ADF) test, ADF^{GLS}_{MAIC}.[9]

In Table 1 we report the results of the $\bar{M}^{GLS}_{MAIC}$ tests and the ADF^{GLS}_{MAIC} test. In all these tests the null hypothesis is that a series is I(1) against the alternative that it is I(0).[10]Our results clearly reject the existence of two unit roots for p_t, at the usual signi...cance levels, while for the ratio c_t/g_t the hypothesis of two unit roots can only be rejected at the 5 % signi...cance level with the ADF^{GLS}_{MAIC} test or at the 10 % with the $\bar{M}^{GLS}_{MAIC}$ tests. The null hypothesis of non-stationarity for the two series in levels can not be rejected in any of the tests applied. Consequently, we can conclude that both variables are I(1).

3.3 Long-term relationship: the intratemporal elasticity of substitution

Once analyzed the order of integration of the series, we are in position to estimate the intratemporal elasticity of substitution, $1/\alpha$.. We estimate the cointegration relationship applying conventional methods to equation (5). Furthermore, we extend our analysis to check whether the long-run

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[8]These tests are the $\bar{M}Z^{GLS}_{\alpha}$, $\bar{M}SB^{GLS}$ and $\bar{M}Z^{GLS}_{t}$.

[9]See Ng and Perron (2001) and Perron and Ng (1996) for a detailed description of these tests.

[10]Note that for the $\bar{M}SB^{GLS}$ test, the null hypothesis is rejected in favour of stationarity when the estimated value is smaller than the critical value.

coefficient estimated is stable over time or exhibits a structural break, leaving the instability to occur at an unknown date.

We estimate equation (5) using the Dynamic Ordinary Least Squares (DOLS)[11] estimation method of Stock and Watson (1993), extended by Shin (1994).[12]Shin (1994) approach is similar to the KPSS[13] tests, which, for the case of cointegration, are implemented in two stages.

Therefore, the first step in our estimation strategy would consist of the estimation of a long run dynamic equation including leads and lags of the explanatory variables in equation (5), i.e. the so-called DOLS regression:

$$p_t = \mu + \delta t + \alpha(c_t/g_t) + \sum_{j=-q}^{q} \varphi_j \Delta (c/g)_{t-j} + \nu_t \qquad (8)$$

where we include a linear trend, t, in order to use Shin's tests.

The second step is the implementation of Shin's tests. These are based on the calculation of two Lagrange statistics (LM): C_μ and C_τ. These are tests for cointegration using the DOLS residuals. The first one tests for deterministic cointegration ($\delta = 0$) whereas the second tests for stochastic cointegration ($\delta \neq 0$).

In Table 2 we report the estimates from the DOLS regression and the results from Shin's tests. We get evidence of stochastic cointegration between p_t and c_t/g_t, being the estimated value $\alpha = 0.72$ with an a priori expected positive sign.[14] This result implies an intratemporal elasticity of substitution $1/\alpha = 1.39$.

As we stated in previous section, we extend previous analysis to check if the intratemporal elasticity of substitution is stable over time or it exhibits any structural break, allowing the instability to occur at a unknown date. We use Gregory and Hansen (1996a, 1996b) approach to test for structural changes in the cointegration relationship.These tests are based on the study of the errors from the long-run regression model of Engle and Granger (1987), in which we include a break in the model with an a priori unknown date, wich would be endogenously determined by the data. There are different alternatives to account for structural breaks in the standard Engle and Granger's cointegration model, although the null hypothesis in all these alternatives is that the series are not cointegrated.

[11] LS estimation of equation (5) might suffer two problems: nuisance parameter dependences due to serial correlation in the residuals and endogeneity bias arising from innovations in the relative price to innovations in consumption.

[12] In order to overcome the problem of the low power of classical tests for cointegration under the presence of persistent roots in the residuals of the cointegration regression, Shin (1994) suggested a new test where the null hypothesis is cointegration.

[13] These tests are called the Kwiatkowski et al. (1992) tests, and assuem the null hypothesis of stationarity.

[14] As we reject the null hypothesis of deterministic cointegration we do not report the value for C_μ in Table 2. This value is available from the authors upon request.

In particular, Gregory and Hansen (1996a, 1996b) proposed four alternative types of regression models to implement the tests: (i) a model with a level shift (the C model); (ii) a model with a linear trend including a level shift (the C/T model); (iii) a model encompassing both a change in the level and in the slope of the coe¢cients of the long term relationship variables (the C/S model), the so-called, "regime shift"; (iv) …nally, an extension of the C/S model (denoted the C/S/T model) that includes a change in the linear trend and so-called "regime and trend shift".[15] To sum up, all these models allow both to detect the existence of cointegration under the presence of time discontinuities, of di¤erent nature, in the long run regression series and to e¢ciently test for the breakpoint date (T_b). The Gregory and Hansen statistics consist in three tests: the modi…ed versions of the $Z_{®}$ and Z_t statistics of Phillips (1987) test, and the augmented Dickey-Fuller test (ADF).[16]

In Table 3 we report Gregory and Hansen (1996a, 1996b) tests statistics. Our results indicate that we can only reject, at 1% signi…cance level, the null of no cointegration in the C/S and the C/S/T models. According to these tests the break point is located between the end of 1973 and the beginning of 1974.

As there is strong evidence of the presence of a structural change in 1974 for the cointegration relationship, we divide our sample in two sub-samples to analyse if the elasticity of intratemporal substitution changes before and after the break. We estimate the cointegration equation (8) for the two sub-samples. These estimates are reported in the last two columns of Table 2. In both cases, we get signi…cant evidence of cointegration between p_t and $c_t=g_t$, being the estimated value for ®, 2:16 and 0:15, respectively. These parameter estimates imply that the value of the intratemporal elasticity of substitution, 1=®, is 0:46 for the …rst sub-sample and 6:66 for the second sub-sample.

3.4 GMM estimation: the intertemporal elasticity of substitution parameter

Using the estimates for ® and ¹ we impose them in the Euler equation (6) to estimate ¯ and °, through GMM.

In table 4 we report the GMM estimation results obtained both from the the full sample and using the pre-break and post break subsamples. In the three cases we …nd that the estimathe for the discount factor is statistically signi…cant and with a value close to the value one would would expect from

[15] See Table 3 for more details of these alternative models.

[16] The lag order of the ADF test (k) is obtain using Ng and Perron (1995) procedure. We take k = 5 as the maximum statrting value and we reduce it progressively up to the point when the t statistic corresponding to the last lag of the …rst di¤erence included in the ADF test is signi…cant at a 5% level, using the normal distribution critical values.

economic theory ($\beta = 1$). Second, for the three cases we get a significant estimate for the parameter γ. The value estimated for this parameter yields an intertemporal elasticity of substitution of 0.74 for the full-sample and of 0.38 and 3.84 for the pre- and post-break sub-samples, respectively. Since we get that $1/\gamma < 1/\alpha$ in every case, we obtain evidence for the Spanish economy that C_t and G_t are Edgeworth-Pareto substitutes.

4 Conclusions

In this paper, we examine the relationship between private and public consumption using Spanish data over the period 1960-2001. We consider a two-good permanent-income model which allows us to estimate both the intratemporal and intertemporal elasticities of substitution between the two types of expenditure.

We estimate the relationship between private and public consumption in Spanish economy in two steps. In the first step, we use cointegration techniques to estimate the intraperiod preference parameter. Thus, we extend the previous analysis addressing the question of whether this long-run relationship is stable over time, or exhibit a structural break allowing the instability to occur at an unknown date. In doing so, we apply recent econometric methodology to detect the structural changes or regime shifts in the cointegration regressions. In the second step, we estimate the intertemporal parameter and the discount factor, using the GMM procedure.Thus, we extend the previous analysis addressing the question of whether this long-run relationship is stable over time, or exhibit a structural break allowing the instability to occur at an unknown date. In doing so, we apply recent econometric methodology to detect the structural changes or regime shifts in the cointegration regressions.

First, our results are consistent with the existence of a long-run relationship between private and public consumption in Spanish economy. Thus, we have been detected the structural changes or regime shifts in the cointegration regression around the time of the oil price shock of 1973/74. Secondly, the estimates both the intratemporal and intertemporal elasticities of substitution between the two types of expenditure (in full-sample and in both sub-samples) suggests that private and public consumption in Spanish economy are Edgeworth-Pareto substitutes, suggesting that an decrease in public consumption will induce an increase in consumption by the private sector. This in turn would support the theoretical insights put forward by Giavazzi and Pagano (1990, 1996) and other authors, who introduced the possibility of "expansionary effects of fiscal contractions" or the "non-keynesian effects" of the fiscal policy.

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Table 1
Ng and Perron[a,b] tests for a unit roots

I(2) vs. I(1)	Case: $p = 0; \bar{c} = -7.0$			
Variable	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	ADF^{GLS}
Δp_t	-15.46^{***}	-2.74^{***}	0.177^{**}	-3.72^{***}
$\Delta c_t/g_t$	-6.37^{*}	-1.76^{*}	0.274^{*}	-2.07^{**}

I(1) vs. I(0)	Case: $p = 1; \bar{c} = -13.5$			
Variable	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	ADF^{GLS}
p_t	-0.15	-0.10	0.683	-0.63
c_t/g_t	-0.93	-0.67	0.720	-1.39

Notes:
[a] A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively.
[b] The MAIC information criteria is used to select the autoregressive truncation lag, k, as proposed in Perron and Ng (1996). The critical values are taken from Ng and Perron (2001), table 1.

Critical values:	Case: $p = 0; \bar{c} = -7.0$			Case: $p = 1; \bar{c} = -13.5$		
	10%	5%	1%	10%	5%	1%
MZ_{α}^{GLS}	-5.7	-8.1	-13.8	-14.2	-17.3	-23.8
MSB^{GLS}	0.275	0.233	0.174	0.185	0.168	0.143
$MZ_{t}^{GLS}; ADF^{GLS}$	-1.62	-1.98	-2.58	-2.62	-2.91	-3.42

Table 2
Stock-Watson-Shin's DOLS[a;b;e] estimation of cointegrating vectors: equation (8)

Parameter Estimates	1960-2001 Full-sample	1960-1974 Pre-break Sub-sample	1975-2001 Post-break Sub-sample
$\hat{\mu}$	-1.55	-3.77	-0.40
	(-6.25)	(-42.9)	(-2.83)
$\hat{\delta}$	0.019	—	0.006
	(7.78)		(3.44)
$\hat{\alpha}$	0.72	2.16	0.15
	(5.30)	(41.6)	(2.30)
$1/\hat{\alpha}$[d]	1.39	0.46	6.66
Test:			
C_{μ}^{c}	—	0.146	—
C_{τ}^{c}	0.079	—	0.093
$\bar{R}^2$	0.95	0.99	0.83
$\hat{\sigma}^2$	0.034	0.020	0.014

Notes:

[a] t-statistics in brackets. Standard Errors are adjusted for long-run variance. The long-run variance of the cointegrating regression residual is estimated using the Barlett window which is approximately equal to $INT\left(T^{1/2}\right)$ as proposed in Newey and West (1987).

[b] We choose $q = INT\left(T^{1/3}\right)$ as proposed in Stock and Watson (1993).

[c] C_{μ} and C_{τ} are LM statistics for cointegration using the DOLS residuals from deterministic and stochastic cointegration, respectively, as proposed in Shin (1994). A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively.

[d] Intratemporal elasticity of substitution parameter.

[e] The critical values are taken from Shin (1994), table 1, from $m = 1$:

Critical values:			
	10%	5%	1%
C_{μ}	0.231	0.314	0.533
C_{τ}	0.097	0.121	0.184

Table 3
Gregory-Hansen[a;b] tests for a single structural change in cointegration relationship: equation (5)

Model	ADF^{*}	T_b	Z_t^{*}	T_b	Z_{α}^{*}	T_b
C[c]	-4.48* (k = 3)	1970	-4.50*	1967	-24.96	1967
C/T[d]	-3.28 (k = 0)	1994	-3.44	1994	-21.26	1994
C/S[c]	-6.31*** (k = 0)	1974	-6.39***	1974	-39.10	1973
C/S/T[d]	-6.78*** (k = 0)	1974	-6.86***	1974	-41.61	1973

Notes:
[a] A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively. The critical values have been obtained from Gregory and Hansen (1996a, 1996b), table 1, m = 1.
[b] The lag order of the tests in brackets.
[c] Deterministic cointegration.
[d] Stochastic cointegration.

ESTIMATED MODELS:
a) "level shift model" (model 2, C):
$p_t = \mu_1 + \mu_2 \varphi_{\tau t} + \alpha_1 (c_t/g_t) + \varepsilon_t$
b) "level shift with trend" (model 3, C/T):
$p_t = \mu_1 + \mu_2 \varphi_{\tau t} + \delta t + \alpha_1 (c_t/g_t) + \varepsilon_t$
c) "regime shift" (model 4, C/S):
$p_t = \mu_1 + \mu_2 \varphi_{\tau t} + \alpha_1 (c_t/g_t) + \alpha_2 (c_t/g_t) \varphi_{\tau t} + \varepsilon_t$
d) "regime and trend shift" (model 5, C/S/T):
$p_t = \mu_1 + \mu_2 \varphi_{\tau t} + \delta_1 t + \delta_2 t \varphi_{\tau t} + \alpha_1 (c_t/g_t) + \alpha_2 (c_t/g_t) \varphi_{\tau t} + \varepsilon_t$

where μ_1, μ_2, α_1, α_2, δ_1 and δ_2 denote the intercept before the shift, the change in the intercept at the time of the shift, the coefficient on the slope of the cointegration relationship, the change in the slope coefficient, and trend coefficient before and after the break, respectively; and $\varphi_{\tau t} = 0$ if $t \leq (\tau T)$, and 1 if $t > (\tau T)$, where the unknown parameter $\tau \in (0,1)$ indicates the (relative) timing of the break point.

Table 4
GMM Estimates of Euler equation (6)[a]

Parameter Estimates	1960-2001 Full-sample[d]	1960-1974 Pre-break[d] Sub-sample	1975-2001 Post-break[e] Sub-sample
$\hat{\beta}$	0.943	0.828	0.976
	(0.014)***	(0.089)***	(0.007)***
$\hat{\gamma}$	1.34	2.63	0.26
	(0.239)***	(1.045)***	(0.143)*
Relationship between private and public consumption:			
$1/\gamma$[b]	0.74	0.38	3.84
$1/\alpha$[c]	1.39	0.46	6.66
$1/\gamma$ T $1/\alpha$	<	<	<
Edgeworth-Pareto sens:	substitutes	substitutes	substitutes

Notes:

[a] Asymptotic standard errors are in parentheses. A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively.

[b] Intertemporal elasticity of substitution parameter.

[c] Intratemporal elasticity of substitution parameter. Source: Table 2.

[d] Instruments set lags: constant, C_t/C_{t-1}, G_t/G_{t-1} and R_{t-1}.

[e] Instruments set lags: constant, C_t/C_{t-1}, G_t/G_{t-1}, C_{t-2}/C_{t-3}, G_{t-2}/G_{t-3}, R_{t-1} and R_{t-2}.