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Estimating the substitutability between private and public consumption: the case of Spain, 1960–2003

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This paper examines the relationship between private and public consumption using Spanish data over the period 1960–2003, using a two-good permanent-income model. We extend previous analysis addressing the question of whether this relationship is stable over time, or exhibits a structural break allowing the instability to occur at an unknown point in time. Our empirical results indicated the existence of a long-run relationship between private and public consumption. We also detect a structural change of regime shift in the cointegration regression around the time of 1973–74. Finally, the estimated intratemporal and intertemporal elasticities of substitution between the two types of expenditure suggest that private and public consumption in Spain are Edgeworth-Pareto substitutes.

I. Introduction

The creation of the Economic and Monetary Union (EMU) by 12 member countries of the European Union (EU) implied that fiscal policy has become the main instrument available to national authorities as stabilization policy. Consequently, issues related to fiscal policy have gained a growing interest in recent years, both in academic and policy environments. In particular, the long-run relationship between private and government consumption has received a great concern for its implications in the process of fiscal consolidation during the 1990s. In the majority of EU countries this consolidation has been planned 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 willingness to substitute its own expenditure for public consumption. For example, if this substitution effect exists, it will offset the effect 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 effectiveness of fiscal policy. If substitutability is empirically accepted, it would support the theoretical insights put forward by Giavazzi and Pagano (1990, 1996), and Barry and Devereux (2003) and more others, who introduced the possibility of 'expansionary effects of fiscal

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contractions' or the 'non-Keynesian effects' of the fiscal policy.[1]

To analyse the relationship between public and private consumption we study the case of Spain for the period 1960–2003. The interest of analysing this case is that Spain is a country traditionally experiencing high budget deficits and which has accomplished an important fiscal consolidation in last years. Such efforts 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 government consumption depends upon whether this variable increases or decreases the marginal utility of private consumption, i.e. whether government and private consumption are Edgeworth-Pareto complements or substitutes. Unfortunately, the empirical works on the substitutability between private and public consumption are still inconclusive due to differences in the econometric methodology, the particular specification of the private agent's utility function and the sample period used.[2]

Previous research for the Spanish economy has found either that private and public consumption are substitutes (Esteve *et al.*, 1997) or that private sector consumption is unaffected by government consumption (Marchante, 1993). In this paper we update these works by incorporating a new model, developed in Amano and Wirjanto (1998). With this model we can estimate both the intratemporal and intertemporal elasticities of substitution for private and public consumption. We can also detect whether the relationship between private and public consumption is stable over time or exhibit structural break allowing the instability to occur at an unknown point in time.

Our empirical results are consistent with the existence of a long-run relationship between private and public consumption. Futhermore, we also detect a structural change or regime shift in the cointegration regression around 1973–74. The estimates intratemporal and intertemporal elasticities of substitution between the two types of expenditure (both in full-sample and in the sub-samples) suggest that private and public consumption for Spain are Edgeworth-Pareto substitutes. These results are in line with the theoretical insights put forward by Giavazzi and Pagano (1990, 1996), and Barry and Devereux (2003) among others, who introduced the possibility of 'expansionary effects of fiscal contractions' or the 'non-Keynesian effects' of fiscal policy.

The rest of the paper is organized as follows. The theoretical framework is presented in Section II. In Section III we describe the data, and present the estimation results. Finally, Section IV summarizes the main conclusions.

II. 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

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

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

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

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

[1] For a survey on the theoretical and empirical literature about non-Keynesian effects of the fiscal policy, and on the European experience for fiscal consolidations see Zaghini (2001).
[2] See Ho (2001) for a recent survey on the literature about the degree of substitutability between government spending and private consumption.
[3] A similar approach is taken in Amano and Wirjanto (1998).
[4] For $1/\alpha = 1$ we assume that $u(C_t, G_t) = C_t^{\phi} G_t^{1-\phi}$.

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, $1/\hat{\alpha}$. Furthermore, we would estimate the intertemporal elasticity of substitution, $1/\gamma$, and the discount factor, β, in a second stage, by imposing the above estimate in the Euler equation.

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 function 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]

III. 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 estimation 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.

Data

We use Spanish annual data for the period 1960–2003 drawn from the Bank of Spain. The variable of interest, the rate C_t/G_t, is obtained using public and private 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.

Unit root tests

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

[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 $sign[U_{CG,t}] = sign[1/\gamma - 1/\alpha]$.
[6] For more details, see Newman (1998).
[7] We thus exploit 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.

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

	$\bar{M}^{GLS}_{MAIC}$ tests			
Variable	$\bar{M}Z^{GLS}_{\alpha}$	$\bar{M}Z^{GLS}_{t}$	$\bar{M}SB^{GLS}$	ADF^{GLS}
I(2) vs. I(1)	Case: $p = 0$, $\bar{c} = -7.0$			
Δp_t	−16.14***	−2.82***	0.175**	−3.82***
$\Delta c_t/g_t$	−6.52*	−1.77*	0.271*	−2.07**
I(1) vs. I(0)	Case: $p = 1$, $\bar{c} = -13.5$			
p_t	−0.24	−0.17	0.702	−0.70
c_t/g_t	−0.94	−0.67	0.710	−1.36

	Case: $p = 0$, $\bar{c} = -7.0$			Case: $p = 1$, $\bar{c} = -13.5$		
Critical values	10%	5%	1%	10%	5%	1%
$\bar{M}Z^{GLS}_{\alpha}$	−5.7	−8.1	−13.8	−14.2	−17.3	−23.8
$\bar{M}SB^{GLS}$	0.275	0.233	0.174	0.185	0.168	0.143
$\bar{M}Z^{GLS}_{t}$, ADF^{GLS}	−1.62	−1.98	−2.58	−2.62	−2.91	−3.42

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.

In general, the majority of the conventional unit root tests (DF and PP types) suffer 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 suffer from severe size distortions when the moving-average polynomial of the first differences 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 modified 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 Modified 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 significance levels, while for the ratio c_t/g_t the hypothesis of two unit roots can only be rejected at the 5% significance 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).

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

[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.

Table 2. Stock-Watson-Shin's DOLS[a, b, e] estimation of cointegrating vectors: Equation 8

Parameter Estimates	Full-sample 1960–2003	Pre-break Sub-sample 1960–1974	Post-break Sub-sample 1975–2003
$\hat{\mu}$	−1.56	−3.77	−0.39
	(−6.09)	(−42.9)	(−2.75)
$\hat{\delta}$	0.019	—	0.006
	(7.57)		(3.36)
$\hat{\alpha}$	0.72	2.16	0.15
	(5.17)	(41.6)	(2.23)
$1/\hat{\alpha}^d$	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
Critical values	10%	5%	1%
C_μ	0.231	0.314	0.533
C_τ	0.097	0.121	0.184

Notes: [a]*t*-statistics are in parentheses. 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 $INT(T^{1/2})$ as proposed in Newey and West (1987).
[b]We choose $q = INT(T^{1/3})$ 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).
[d]Intratemporal elasticity of substitution parameter.
[e]The critical values are taken from Shin (1994), Table 1, from $m = 1$.

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's (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:

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

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 $1/\hat{\alpha} = 1.39$.

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

[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 assume 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.

Table 3. Gregory-Hansen[a, b] tests for a single structural change in the cointegration relationship: Equation 5

Model	ADF^*	T_b	Z_t^*	T_b	Z_α^*	T_b
C[c]	−4.63** ($k = 3$)	1970	−4.60*	1967	−26.10	1967
C/T[d]	−3.34 ($k = 0$)	1994	−3.48	1994	−21.19	1994
C/S[c]	−6.52*** ($k = 0$)	1974	−6.60***	1974	−40.97	1974
C/S/T[d]	−6.94*** ($k = 0$)	1974	−7.02***	1974	−43.40	1974

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 (k) of the tests are in parentheses.
[c]Deterministic cointegration.
[d]Stochastic cointegration.

ESTIMATED MODELS:
(a) "level shift model" (model C):
$p_t = \mu_1 + \mu_2\varphi_{\tau t} + \alpha_1(c_{t/}g_t) + \epsilon_t$
(b) "level shift with trend" (model C/T):
$p_t = \mu_1 + \mu_2\varphi_{\tau t} + \delta t + \alpha_1(c_t/g_t) + \epsilon_t$
(c) "regime shift" (model C/S):
$p_t = \mu_1 + \mu_2\varphi_{\tau t} + \alpha_1(c_t/g_t) + \alpha_2(c_t/g_t)\varphi_{\tau t} + \epsilon_t$
(d) "regime and trend shift" (model C/S/T):
$p_t = \mu_1 + \mu_2\varphi_{\tau t} + \delta_1 t + \delta_2 t\varphi_{\tau t} + \alpha_1(c_t/g_t) + \alpha_2(c_t/g_t)\varphi_{\tau t} + \epsilon_t$
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; $\varphi_{\tau t}$ is 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.

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, which 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.

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 coefficients of the long term relationship variables (the C/S model), the so-called, "regime shift"; (iv) finally, 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 different nature, in the long run regression series and to efficiently test for the breakpoint date (T_b). The Gregory and Hansen statistics consist in three tests: the modified versions of the Z_α 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% significance 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 subsamples 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

[15] See Table 3 for more details of these alternative models.
[16] We obtain the lag order of the *ADF* test (k) using Ng and Perron (1995) procedure. We take $k = 5$ as the maximum starting value and we reduce it progressively up to the point when the t statistic corresponding to the last lag of the first difference included in the *ADF* test is significant at a 5% level, using the normal distribution critical values.

Table 4. GMM Estimates of the Euler Equation 6[a]

Parameter Estimates	Full-sample[d] 1960–2003	Pre-break[d] Sub-sample 1960–1974	Post-break[e] Sub-sample 1975–2003
$\hat{\beta}$	0.973	0.828	0.976
	(0.009)***	(0.089)***	(0.007)***
$\hat{\gamma}$	1.07	2.63	0.27
	(0.119)***	(1.045)***	(0.182)*
Relationship between private and public consumption			
$1/\hat{\gamma}^{b}$	0.93	0.38	3.70
$1/\hat{\alpha}^{c}$	1.39	0.46	6.66
$1/\hat{\gamma} \gtreqless 1/\hat{\alpha}$	$<$	$<$	$<$
Edgeworth-Pareto sens	*substitutes*	*substitutes*	*substitutes*

Notes: [a]Asymptotic standard errors are in parentheses. A * and *** denote significance at the 10% and 1% levels, respectively.
[b]Intertemporal elasticity of substitution parameter.
[c]Intratemporal elasticity of substitution parameter. Source: Table 2.
[d]Instrument set: constant, C_t/C_{t-1}, G_t/G_{t-1} and R_{t-1}.
[e]Instrument set: 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}.

in the last two columns of Table 2. In both cases, we get significant evidence of cointegration between p_t and c_t/g_t, being the estimated value for $\hat{\alpha}$, 2.16 and 0.15, respectively. These parameter estimates imply that the value of the intratemporal elasticity of substitution, $1/\hat{\alpha}$, is 0.46 for the first sub-sample and 6.66 for the second sub-sample.

GMM estimation: the intertemporal elasticity of substitution parameter

Using the estimates for α and μ we impose them in the Euler Equation 6 to estimate β and γ, through GMM.

In Table 4 we report the GMM estimation results obtained both from the full sample and using the pre-break and post break subsamples. In the three cases we find that the estimate for the discount factor is statistically significant and with a value close to the value one would expect from 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 0.93 for the full-sample and of 0.38 and 3.70 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.

IV. Conclusions

In this paper, we examine the relationship between private and public consumption for Spain over the period 1960–2003. 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.

To estimate the relationship between private and public consumption we proceed in two steps. In the first step, we use cointegration techniques to estimate the intraperiod preference parameter. We also extend previous analysis to address 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 structural changes or regime shifts in cointegration regressions. In the second step, we estimate the intertemporal parameter and the discount factor, using the GMM procedure.

Our results indicate the existence of a long-run relationship between private and public consumption in the Spanish economy. Futhermore, we also detect a structural change or regime shift in the cointegration regression around the 1973–74. According to the estimates of both the intratemporal and intertemporal elasticities of substitution between the two types of expenditure we get that private and public consumption are Edgeworth-Pareto substitutes for Spain. This result supports the theoretical insights put forward by Giavazzi and Pagano (1990, 1996), and Barry and Devereux (2003) among others, who introduced the possibility of 'expansionary effects of fiscal contractions' or the 'non-Keynesian effects' of the fiscal policy.

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