Economic Modelling 27 (2010) 782–787

Contents lists available at ScienceDirect

Economic Modelling

journal homepage: www.elsevier.com/locate/ecmod

A panel cointegration approach to estimating substitution elasticities in consumption

Monica Auteri [a], Mauro Costantini [b,*]

[a] Dipartimento di Istituzioni Pubbliche, Economia e Società - Università Roma Tre, Italy
[b] Department of Economics, University of Vienna, Austria

ARTICLE INFO

Article history:
Accepted 25 January 2010

JEL Classification:
C22
C23
E21
E62
H5

Keywords:
Cointegration
Government expenditure
Intratemporal substitution
Macroeconomic aspects of public finance
Panel
Unit root

ABSTRACT

This paper investigates the relationship between government spending and private consumption. The general framework is a cointegration approach of Ogaki (1992) used to estimate the intratemporal elasticity of substitution between government and private consumption in a panel of 15 European countries. Recently developed non-stationary panel methodologies that assume cross-section dependence are applied. Results indicate an Edgeworth substitutability between private and public spending.

1. Introduction

The relationship between public and private consumption has come to play an important role for theoretical and empirical models that form the mainstream paradigm in contemporary macroeconomics and public finance. The issue whether public and private consumption are complementary or substitute goods is not without merit. To give an example that underscores the need for an empirical answer, Djajic (1987) shows that for a thorough understanding of the fiscal policy effects of government spending it is of crucial importance to know whether the public and private consumption goods are independent, substitutes or complements.

Governments spend a large part of their budget on goods and services that can be privately produced and are likely to affect consumer's utility in a similar way to private consumption, therefore a decision on appropriate fiscal policy needs to take into account the substitutability between private and public consumption. If the private sector derives utility from public provided goods and services and perceives private and government consumption as substitutes, a rise in government spending determines a fall in private consumption, due to crowding-out. The higher the substitutability between private and government consumption the larger the crowding-out effect on private consumption. On the other hand if private and government consumption are complements, the effectiveness of fiscal policy in stimulating aggregate demand will be amplified by an increase in private consumption.

Despite the relevance of the issue, we know surprisingly little about the effects of fiscal policy on private consumption and on economic activities in general. The question of whether private consumption and public spending are complements or substitutes has been examined by several studies such as those by Aschauer (1985), Karras (1994), Ni (1995), Amano and Wirjanto (1998), Van Dalen (1999) and Okubo (2003). Some recent contributions in panel data are from Nieh and Ho (2006) and Kwan (2006). These studies all use a partial equilibrium approach based on Euler equations to estimate the degree of substitutability between private and public spending. The empirical results, however, are mixed and inconclusive due to differences in the econometric methodology. It remains that the existing literature does not seem to provide much guidance as to the magnitude of the elasticity of substitution between private and public spending. This paper attempts to revisit this issue by contributing to the literature in two important aspects. First, the effectiveness of fiscal expansion is evaluated as the intratemporal elasticity of substitution between government spending and private consumption in a panel of 15 European countries over the 1970–2007 period.

* Corresponding author. Department of Economics, University of Vienna, BWZ - Bruenner Strasse 72, A-1210 Vienna, Austria. Tel.: +43 1 4277 37478; fax: +43 1 4277 37498.
E-mail addresses: mauteri@uniroma3.it (M. Auteri), mauro.costantini@univie.ac.at (M. Costantini).

0264-9993/$ – see front matter © 2010 Elsevier B.V. All rights reserved.
doi:10.1016/j.econmod.2010.01.021

In the current institutional setting, even if in the EU there is no single fiscal policy, "high degree of fiscal synchronization that exists within the EU [...] cannot be ignored without simultaneously incurring a cost in efficiency" (Prohl and Westerlund, 2009). Since fiscal policies tend to be relatively homogeneous and preference parameters are similar, there is a need of an approach that takes full account of the panel structure of the data. One of the most important attractions of panel data is the ability to pool the long-run information regarding the cointegration parameters. In this paper, we use a cointegration approach of Ogaki (1992) and Ogaki and Park (1997) to estimate the preference parameters that govern the relationship between government and private consumption. Cointegration is applied to exploit the long-run restriction imposed by the intraperiod first-order condition. This method can directly evaluate the intratemporal elasticity of substitution from the data that is shown to be robust to a number of economic factors.

Second, we take into account cross-sectional dependence through factor models to mirror the behavior of fiscal authorities in the context of the EU Treaty, the setting up of the convergence criteria, and the Stability and Growth Pact, SGP. Within the EU framework, fiscal discipline is determinant for the implementation of the common monetary policy. The political and economic cohesion of monetary union requires fiscal support across regions and in EMU, and the existence of sound fiscal policies is seen as a necessary objective for individual countries to pursue.[1] In this context there are possible cross-country dependencies that can be envisaged not only in the run-up to EMU but also, for example, via integrated financial markets.[2] In addition, policy choices are often affected by neighboring countries via externalities among jurisdictions. An example of these externalities is the amount of public investments in infrastructures in a country, such as roads, airports and rail-tracks. The benefits of these investments spill over in neighboring countries, and therefore affect the level of investments in the latter countries. Other interdependencies that occur with geographically close countries are due to tax competition (countries compete with their neighbor in order to attract tax base since corporate taxes are more sensitive to those of closer countries) and to the tendency of citizens to evaluate the performances of their policy makers by comparing the same policy choices taken by the neighboring countries (the so called "yardstick" competition).[3]

The organization of the remainder of this paper is as follows. Section 2 lays out the model and the intratemporal cointegration relationship implied is derived and discussed. Section 3 describes the econometric methodology and presents results. Section 4 concludes.

2. The model

The empirical analysis in this paper follows Amano and Wirjanto (1997, 1998) who make use of the cointegration approach of Ogaki (1992) and Ogaki and Park (1997) to estimate the preference parameters that govern the relationship between government and private consumption.[4] In this model, a representative agent maximizes the expected lifetime consumption utility function expressed as:

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

where C_t is real private consumption at time t and G_t is real government expenditures in period t. Eq. (1) is subject to a lifetime budget constraint in a complete market at period t. E_t is the expectations operator based on period t information, and $0 \le \beta \le 1$ is a discount factor.

Consider the addilog utility function:

$$U(C_t, G_t) = \left(\frac{C_t^{1-\alpha}}{1-\alpha}\right)\Lambda_{C_t} + K\left(\frac{G_t^{1-\nu}}{1-\nu}\right)\Lambda_{G_t} \tag{2}$$

where α and ν are curvature parameters with α and $\nu \ge 0$, K is a scaling factor, and Λ_{C_t} and Λ_{G_t} represent random preference shocks associated with private and public consumption, respectively. By allowing for these shocks, we avoid the assertion of Garber and King (1983) that the presence of random preference shocks can often yield misleading results. Hence the representative consumer maximizes the intraperiod utility of Eq. (2) subject to the intratemporal budget constraint below:

$$P_{c,t}C_t + P_{g,t}G_t = M_t \tag{3}$$

where $P_{c,t}$ and $P_{g,t}$ are the prices of private consumption and public consumption at time t and M_t is the total consumption expenditure at time t. We assume that the sequences of random preference shocks Λ_{C_t} and Λ_{G_t} are stationary or $I(0)$ processes. An intratemporal (or static) first-order necessary condition of the above-mentioned problem states that the relative purchase price of government to private consumption is equal to the marginal rate of substitution based on the purchase of the two types of goods, i.e.:

$$\frac{C_t^{-\alpha}\Lambda_{C_t}}{P_{c,t}} = \lambda_t \tag{4}$$

$$\frac{KG_t^{-\nu}\Lambda_{G_t}}{P_{g,t}} = \lambda_t \tag{5}$$

then, combining Eqs. (4) and (5), we obtain:

$$\frac{P_{g,t}}{P_{c,t}} = \frac{KG_t^{-\nu}}{C_t^{-\alpha}}\frac{\Lambda_{G_t}}{\Lambda_{C_t}} \tag{6}$$

Taking logarithms on both sides of Eq. (6) and rearranging yields:

$$lnG_t = \mu + \left(\frac{\alpha}{\nu}\right)lnC_t - \left(\frac{1}{\nu}\right)lnP_t + \varepsilon_t \tag{7}$$

where $P_t = P_{g,t}/P_{c,t}$, $\mu = \left(\frac{1}{\nu}\right)lnK$ and ε_t collects the remaining terms and represents a stationary process with zero mean; Eq. (7) implies that $\ln P_t$, $\ln G_t$ and $\ln C_t$ are cointegrated with cointegrating vector $[1, \nu, -\alpha]$.[5] We assume that all cross-sectional units have similar preference parameters, hence we estimate the panel form of Eq. (7) as specified below:

$$g_{i,t} = \mu_i + \left(\frac{\alpha}{\nu}\right)c_{i,t} - \left(\frac{1}{\nu}\right)p_{i,t} + \varepsilon_{i,t}. \tag{8}$$

The lower case letters denote variables in logarithmic form. Eq. (8) has important implications. First, government spending affects the private consumption by ν/α, hence, the effectiveness of a fiscal expansion is determined by the magnitude of parameter estimate.

[1] Nowadays countries have to comply with the budgetary requirements of EU, by avoiding excessive deficits, keeping debt levels below the 60% of GDP reference value, and respecting the requirements of the SGP.

[2] Indeed, with cross-country spillovers in government bond markets, interest rates comovements inside the EU became also more noticeable, especially after the completion of the single EU15 capital market in 1994.

[3] In the European framework, the "yardstick" competition has been empirically investigated among others by Redoano (2003). He found empirical support to the idea that EU governments influence each others in determining their fiscal choices.

[4] These studies use USA data with different frequency and span. In particular, Amano and Wirjanto (1997, 1998) and Ogaki and Park (1997) use quarterly data over the period 1953–1994 and 1947–1983 respectively, while Ogaki (1992) uses annual data over the period 1929–1988.

[5] It should be noted that both G and C are treated as choice variables for a representative agent that maximizes his (or her) utility by optimally consuming both private and government goods.

Table 1
Individual unit root and cointegration results.

Country	g_t				c_t				p_t				Coint. ADF
	$DF-GLS^{\mu}$	$DF-GLS^{\tau}$	s_α	s_β	$DF-GLS^{\mu}$	$DF-GLS^{\tau}$	s_α	s_β	$DF-GLS^{\mu}$	$DF-GLS^{\tau}$	s_α	s_β	
Austria	0.298(3)	−2.715(3)	3.589(1)	0.345(1)	0.976(3)	−1.983(1)	3.505(1)	0.256(1)	−1.533(1)	−1.600(1)	1.787(1)	0.297(1)	−4.472
Belgium	1.246(2)	−1.831(1)	3.412(1)	0.318(2)	0.566(1)	−2.329(1)	2.506(1)	0.161(1)	−0.954(1)	−2.278(1)	1.916(1)	0.220(1)	−5.824
Denmark	0.262(2)	−1.098(2)	4.265(1)	0.450(1)	1.546(1)	−1.807(1)	3.369(2)	0.407(2)	−1.259(1)	−2.236(1)	0.497(1)	0.197(1)	−3.751
Finland	−0.017(1)	−1.156(1)	2.235(1)	0.259(1)	−0.918(1)	−2.498(1)	3.238(2)	0.232(2)	−0.947(2)	−0.917(2)	2.500(2)	0.372(2)	−3.660
France	0.264(3)	−1.516(1)	3.074(1)	0.310(1)	0.872(1)	−2.181(1)	3.272(2)	0.156(2)	−1.421(2)	−2.053(2)	0.599(1)	0.197(1)	−4.564
Germany	−0.109(2)	−1.849(2)	2.791(1)	0.293(1)	−0.649(1)	−1.955(1)	1.175(1)	0.235(1)	−1.251(3)	−1.415(3)	2.163(1)	0.239(1)	−3.951
Greece	−0.451(2)	−2.007(2)	1.789(2)	0.313(2)	0.764(1)	−2.021(1)	2.976(2)	0.313(2)	0.485(1)	−2.316(1)	2.746(1)	0.314(1)	−4.721
Ireland	0.718(1)	−1.699(1)	3.122(2)	0.421(2)	0.797(1)	−1.568(1)	3.115(2)	0.266(2)	1.837(1)	−1.153(1)	3.271(1)	0.319(1)	−4.511
Italy	0.621(1)	−1.961(1)	1.591(1)	0.268(1)	0.873(1)	−2.057(1)	3.185(1)	0.239(1)	−1.583(1)	−2.602(1)	1.555(1)	0.241(1)	−3.871
Luxemb.	−1.084(1)	−1.733(1)	1.789(1)	0.271(1)	−0.351(1)	−2.001(1)	4.622(1)	0.211(1)	−0.881(1)	−2.028(1)	2.846(1)	0.581(1)	−3.187
Netherl.	1.398(1)	−2.355(1)	3.706(1)	0.180(1)	−0.457(1)	−1.211(1)	3.707(1)	0.197(1)	−1.770(1)	−2.031(1)	0.495(1)	0.188(1)	−3.101
Portugal	0.431(1)	−0.322(1)	1.912(1)	0.178(1)	0.749(1)	−2.199(1)	3.576(1)	0.335(1)	−1.763(2)	−2.485(2)	1.373(1)	0.251(1)	−4.521
Spain	−0.095(2)	−3.011(2)	3.613(1)	0.259(1)	0.176(1)	−2.026(1)	2.485(2)	0.219(2)	−1.707(1)	−2.688(1)	0.413(1)	0.127 (1)	−4.112
Sweden	−0.256(1)	−1.276(1)	0.461(1)	0.156(1)	0.108(1)	−1.998(1)	2.626(1)	0.411(1)	−0.855(1)	−1.895(1)	0.987(1)	0.236 (1)	−5.123
UK	1.144(1)	−1.629(1)	3.244(1)	0.528(1)	0.476(1)	−2.699(1)	3.680(1)	0.371(1)	−1.439(2)	−0.459(2)	2.794(1)	0.530(1)	−3.980

Notes: $DF-GLS^{\mu}$ and $DF-GLS^{\tau}$ indicate the modified Dickey-Fuller test developed by Elliott et al. (1996) for the model with constant and constant and trend respectively. Numbers in parentheses for the DF-GLS test denote the number of lags selected using the modified AIC criteria (MAIC) with the maximum set equal to $p_{max}=4$. The critical values for the $DF-GLS^{\mu}$ test are −1.61, −1.95 and −2.60 at 10%, 5% and 1% and for $DF-GLS^{\tau}$ are −2.89,−3.19 and −3.77 at 10%, 5% and 1%. s_α and s_β refer to the tests of Leybourne and McCabe (1994) for level and trend model. Numbers in parentheses for these tests indicates the AIC lag order estimated of the ARIMA(p,1,1). Lag orders are constrained to lie between 0 and 4. For s_α and s_β tests, the approximate asymptotic critical values at 10%, 5% and 1% significance level for the level model are, respectively, 0.347, 0.463, and 0.739; for the trend model, they are 0.119, 0.146, and 0.216. The cointegration tests are based on a regression with a constant term. The 5% and 10% critical values for the ADF statistic for cointegration are −3.365 and −3.065. These appear in Phillips and Ouliaris (1990).

Then, ν/α is defined as the intratemporal elasticity of substitution between government and private spending and reflects the willingness of an individual to substitute between private and government consumption within a given period. Second, the parameter $1/\alpha$ is the intertemporal elasticity of substitution for private consumption, or the rate at which consumers substitute consumption between periods in response to relative price changes across periods. The elasticity of intertemporal substitutability determines the responsiveness of consumption over time to changes in the relative price of present and future consumption. Hence, our empirical analysis focuses on estimating parameters ν and α to calculate these elasticities within a panel data set in order to determine whether private and government consumption are substitutes or complements.[6] According to Amano and Wirjanto (1997), Eq. (8) has three testable implications, summarized by Nieh and Ho (2006) as follows:

1. If $\left(\frac{\nu}{a}\right)<\left(\frac{1}{\alpha}\right)$, the intratemporal elasticity of substitution is smaller than the intertemporal elasticity of substitution and then C and G are Edgeworth-Pareto complements;
2. If $\left(\frac{\nu}{a}\right)>\left(\frac{1}{\alpha}\right)$, the intratemporal elasticity of substitution dominates over the intertemporal elasticity of substitution and then C and G are Edgeworth-Pareto substitutes; and
3. If $\left(\frac{\nu}{a}\right)=\left(\frac{1}{\alpha}\right)$, then C and G are Edgeworth independent, or unrelated.

3. Econometric methodology and results

In order to test the effectiveness of fiscal expansion, the intratemporal elasticity of substitution between government spending and private consumption is estimated using a cointegration approach. Preliminary univariate analysis is carried out in order to compare time-series evidence for each single country to panel data estimates. Results for panel data under cross-section dependence hypothesis are then presented.[7] The data set consists of annual data of 15 European countries over the period 1970–2007.[8] Data is taken from OECD Statistical Compendium (2007). The per-capita consumption series is obtained by dividing personal real consumption of nondurable goods and services by the total population ages 16 and over. The per-capita government expenditure series is measured as the ratio of national real government purchases of goods and services to the total population ages 16 and over. The relative price measure is simply the ratio of public to private expenditure prices where government and private prices are respectively the deflator of government and private consumption.

3.1. Time series properties

As a first step, we explore the nonstationary properties of the data using the ADF-GLS test of Elliott et al. (1996) and the Leybourne and McCabe (1994) tests, hereafter LMC. The ADF-GLS test tests the null of unit root whereas the LMC test tests the null of stationarity. Using these two tests, one can perform a confirmatory analysis on the stationarity properties of the series.[9] If both tests fail to reject the null hypotheses, one can only conclude that data is not informative enough on the stationarity of the series. On the other hand, if the LMC test rejects the null and at the same time the ADF-GLS test fails to reject the null, then both tests support the same conclusion, that is, the series in question is a unit root process.

In Table 1, country-by-country results of the ADF-GL and LMC tests are shown. Reading across the rows the presence of unit root is evident in each countries. Since nonstationarity is found for all variables, we proceed to test for cointegration using the ADF residuals-based test. Results are reported in Table 1 (last column). Evidence of a cointegrating relationship between government spending, consumptions, and price-ratio variables is found for all countries since the null of no-cointegration is always rejected. Next, the cointegrating relationship is estimated with the fully modified OLS (FMOLS) estimator of Phillips and Hansen (1990). The estimates results and the elasticities of substitution are reported in Table 2.

[6] (Ogaki and Park, 1997) show that this cointegration approach allows for non-orthogonal but stationary multiplicative measurement error, the presence of liquidity constraints, and a general form of time separability in preferences. The latter result, however, holds only under the restrictive assumption of additive separability between the two goods. In the case of non-separability, the cointegration approach is not robust to time non-separability except in a few special cases. One such case is given by the intraperiodutility function of the form $U(C_t,G_t)=\Lambda_t(C_t^*)^{\alpha}(G_t^*)^{\nu}$, where $\Lambda_t=(\Lambda_{C_t},\Lambda_{G_t})$ and C_t^* and G_t^* are service flows from purchases of private and public goods, respectively.

[7] Results under cross-section independence hypothesis are available upon request.

[8] Countries are: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain and UK.

[9] We thank one of the referees for suggesting the use of these two most powerful tests.

Table 2
Individual results FMOLS Phillips and Hansen (1990) estimator.

Country	p_t		c_t		Elasticities	
	Estimate	Std. error	Estimate	Std. error	ν/α	$1/\alpha$
Austria	1.579	0.892	1.487	0.698	0.672	−1.062
Belgium	−0.561	0.159	1.387	0.503	0.721	0.404
Denmark	1.305	0.764	1.713	0.756	0.584	−0.762
Finland	−0.598	0.293	1.321	0.623	0.757	0.453
France	−0.419	0.210	1.645	0.350	0.608	0.255
Germany	−0.239	0.114	1.538	0.546	0.650	0.155
Greece	1.487	0.872	1.581	0.501	0.633	−0.941
Ireland	−0.394	0.184	1.321	0.427	0.757	0.298
Italy	−0.522	0.129	1.492	0.521	0.670	0.350
Luxembourg	1.577	0.859	1.503	0.436	0.665	−1.049
Netherlands	−0.455	0.212	1.399	0.541	0.715	0.325
Portugal	1.641	0.872	1.412	0.612	0.708	−1.162
Spain	−0.452	0.101	1.732	0.521	0.577	0.261
Sweden	−0.418	0.201	1.703	0.389	0.587	0.245
UK	−0.277	0.114	1.544	0.385	0.648	0.179

The values of ν/α are always larger than $1/\alpha$, therefore government and private spending are Edgeworth-Pareto substitutes. This would imply that a fiscal contraction may induce substantial expansion in private consumption, thereby offsetting, or even outweighing, the negative impact of the fiscal contraction on aggregate demand. It is noteworthy that the estimate of $1/\alpha$ varies widely across countries and only few of them do not have the sign predicted by the theory.[10]

3.2. Panel data analysis

In the previous subsection, the long-run elasticities of substitution between government spending, private consumption and price-ratio are estimated using time-series approach. However, it is well known that the time-series unit root and cointegration tests have limited power in finite samples.[11] The use of panel data is generally considered as a means of generating more powerful tests, even if a higher power can be obtained at cost of the imposition of the poolability assumption. This is not the case here since the cross-section dependence hypothesis, under which our analysis is performed, allows for additional variation in the cross-section dimension and assures more precise estimates with respect to the estimation results at the level of single-country. Thus, using pool data to estimate the intratemporal elasticity of substitution in a context characterized by relatively homogeneous fiscal policies and similar preference parameters, there are not significant costs involved and cross-correlation assures more precise estimates.

Given the high degree of economic interdependence in the EU, important externalities from national fiscal policies exist. In fact, if it is true that States usually act interdependently when they take their policy choices both with respect to expenditures and taxes, it would also be true that tax competition occurs with geographically close countries, as well as "yardstick" competition, especially with respect to those expenditures that are more directly comparable, i.e. education and health. In addition countries may be affected by their "neighbors" via externalities among jurisdictions whose benefits spill over in neighboring countries and therefore affect the level of investments in the latter countries making policy choices not independent. Therefore in the EU, cross-country spillovers can be envisaged in the presence of similar policy measures which justifies the hypothesis of cross-country dependence.

As in the previous subsection, we first investigate unit root properties of the variables and the existence of a cointegration relationship between them.

[10] This occurs for Austria, Denmark, Germany, Luxembourg and Portugal. Possible explanations for the negative estimates of these elasticities can be found in Ogaki and Reinhart (1998).

[11] See Campbell and Perron (1991) on the power of the univariate tests.

The preference parameters are then estimated using the FMOLS estimator under the hypothesis of cross-section dependence. To get a feeling of the size of the cross-sectional dependence problem in the data, we compute the long-run cross-sectional correlation matrix of the OLS residuals obtained from Eq. (8). Table 3 reports the correlation matrix. Results show that all correlations lie between 0.08 and 0.87, with an overall average of 0.39, suggesting that the independence assumption is clearly violated.

In addition to the cross-correlation residuals matrix, we report results for cross-section dependence assumption in the data using the CD test developed by Pesaran (2004). The CD test is based on a simple average of all pair-wise correlation coefficients of the OLS residuals from the individual regressions in the panel:

$$y_{it} = \alpha_i + \beta_i' x_{it} + u_{it} \text{ for } i = 1, 2, ..., N; t = 1, 2, ..., T, \tag{9}$$

where i indexes the cross-section dimension and t the time series dimension, x_{it} is a $k \times 1$ vector of observed time-varying regressors. The sample estimate of the pair-wise correlation of the residual is:

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^{T} e_{it} e_{jt}}{\left(\sum_{t=1}^{T} e_{it}^2\right)^{1/2} \left(\sum_{t=1}^{T} e_{jt}^2\right)^{1/2}} \tag{10}$$

where e_{it} is the OLS estimate of u_{it} defined by

$$e_{it} = y_{it} - \hat{\alpha}_i - \hat{\beta}_i' x_{it}. \tag{11}$$

The test proposed by Pesaran (2004) is:

$$\mathrm{CD} = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij} \right). \tag{12}$$

where accent $\hat{\rho}_{ij}$ is the pair-wise correlation. In our analysis, we first compute OLS residuals from the regression of any variable of interest on an intercept, a linear trend and a lagged dependent variable for each country i. Using these residuals, we compute the CD statistics. Clear evidence of cross-section dependence for all variables is founded. The results are 3.79 (0.028), 56.65 (0.00) and 11.98 (0.00) for government spending, private consumption and price-ratio variables respectively (p-values are in parentheses).

Then, the presence of a unit root in our series of interest is tested with the so-called PANIC approach (Panel Analysis of Non-stationary in Idiosyncratic and Common components) of Bai and Ng (2004). The CADF test of Pesaran (2007) is also considered in order to provide robustness analysis.[12] The Bai and Ng (2004) approach tests for the presence of a unit root in the common factors and idiosyncratic components separately. It determines then, if non-stationarity comes from the common or from the idiosyncratic components. The PANIC approach is applied to the variables $g_{i,t}$, $c_{i,t}$ and $p_{i,t}$, but first, using Bai and Ng's *IC3* criterion, the number of common factors is selected. The results of the unit root tests for the common factors are reported in Table 4. There is only one common factor for all variables. In this case, Bai and Ng (2004) suggest using a standard Augmented Dickey-Fuller (ADF) test to check for stationarity in the following model:

$$\Delta F_{1,t} = c + \gamma_{i,0} \hat{F}_{1,t-1} + \ldots\ldots + \gamma_{i,p} \hat{F}_{1,t-p} + \nu_{it} \tag{13}$$

where F_t indicates an $r \times 1$ vector of common factors. The ADF tests results for the extracted common factor show evidence of unit root in all variables.

[12] The use of the CADF test has been suggested by a referee. Specifically, we apply $\overline{CADF}^*$ which is the truncated version of $\overline{CADF}$. Pesaran (2007) suggests the use of the truncated version for computational reasons. See footnote to Table 4.

Table 3
Estimated long-run cross-sectional correlations.

No	Country	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15
1	Austria	1.00														
2	Belgium	0.34	1.00													
3	Denmark	0.43	0.31	1.00												
4	Finland	0.35	0.42	0.29	1.00											
5	France	0.28	0.47	0.12	0.09	1.00										
6	Germany	0.78	0.82	0.73	0.21	0.84	1.00									
7	Greece	0.23	0.32	0.12	0.09	0.24	0.18	1.00								
8	Ireland	0.21	0.23	0.12	0.29	0.45	0.29	0.14	1.00							
9	Italy	0.48	0.53	0.33	0.43	0.77	0.68	0.42	0.64	1.00						
10	Luxembourg	0.39	0.08	0.43	0.53	0.58	0.49	0.12	0.47	0.69	1.00					
11	Netherlands	0.15	0.75	0.45	0.29	0.45	0.78	0.51	0.13	0.29	0.39	1.00				
12	Portugal	0.19	0.27	0.57	0.29	0.61	0.58	0.45	0.29	0.78	0.28	0.39	1.00			
13	Spain	0.09	0.23	0.34	0.29	0.68	0.39	0.82	0.38	0.87	0.29	0.39	0.87	1.00		
14	Sweden	0.44	0.22	0.55	0.65	0.25	0.49	0.21	0.67	0.43	0.41	0.76	0.17	0.09	1.00	
15	UK	0.28	0.38	0.43	0.34	0.41	0.54	0.27	0.86	0.41	0.32	0.38	0.29	0.32	0.45	1.00

To test the stationarity of the idiosyncratic component, Bai and Ng (2004) propose to pool individual ADF t-statistics with de-factored estimated components $\hat{e}_{it}$ in the model with no deterministic trend:

$$\Delta e_{i,t} = \delta_{i,0}\,\hat{e}_{i,t-1} + \sum_{j=1}^{p} \delta_{i,j}\Delta\,\hat{e}_{i,t-j} + \mu_{i,t}. \tag{14}$$

The pooled tests are based on Fisher-type statistics defined as in Maddala and Wu (1999)) and in Choi (2001). Let $P_{\hat{e}}^{c}(i)$ be the p-value of the ADF t-statistics for the i-th cross-section unit, $ADF_{\hat{e}}^{c}(i)$, then the standardized Choi's type statistics is:

$$Z_{\hat{e}}^{c} = \frac{-2\sum_{i=1}^{N} logP_{\hat{e}}^{c}(i) - 2N}{\sqrt{4N}} \tag{15}$$

The statistics Eq. (15) converge for $(N, T\rightarrow\infty)$ to a standard normal distribution. In our empirical analysis, we use the Fisher-type statistic defined as in Choi (2001). Table 4 reports the unit root results for the $\overline{CADF}*$, $ADF_{\hat{F}}^{c}$ and $Z_{\hat{e}}^{c}$ tests. Strong evidence in favor of a unit root process is found for all variables.

To investigate the existence of a cointegrating relationship between government spending, private consumption and price-ratio, our testing strategy builds on the approach briefly sketched in Gengenbach et al. (2006) which involves two steps. First, a PANIC analysis, as proposed by Bai and Ng (2004), is applied. Then in the second step:

(1) If the common factors are $I(1)$, but the idiosyncratic components are $I(0)$, the non-stationarity is entirely driven by a reduced number of common stochastic trends, and this implies the case of cross member cointegration. $X_{i,t}$ and $Y_{i,t}$ are cointegrated only if the common factors of the variables cointegrate.

(2) If both, the common factors and idiosyncratic components, are $I(1)$, defactored data is used ($X_{i,t}$ and $Y_{i,t}$ are defactored separately) to test cointegration with Pedroni (2001, 2004) approach.[13]

Table 4
Panel unit root results.

Variables	$\overline{CADF}$	$BN_{ADF_{\hat{F}}}$	$BN_{Z_{\hat{e}}}$
$g_{i,t}$	(0.843) −1.526	(0.470) −1.464	(0.651) −1.084
$c_{i,t}$	(0.980) −0.359	(0.530) −1.299	(0.800) −0.826
$p_{i,t}$	(0.160) −2.011	(0.710) −1.089	(0.532) −1.403

Notes: The number of the common factors estimated according to $IC3$ Criteria is equal to 1. The maximum number of factors is fixed to 4. $BN_{ADF_{\hat{F}}}$ and $BN_{Z_{\hat{e}}}$ denote (Bai and Ng, 2004) unit root tests on common factor and idiosyncratic component respectively. p-values are in brackets. $\overline{CADF}*$ denotes the truncated version of the $\overline{CADF} = N^{-1}\sum_{i=1}^{N} CADF_i$ test, where $CADF_i$ is the individual t-statistics of b_i of the following cross-sectionally augmented equation $\Delta Y_{it} = \alpha_i + b_i Y_{it-1} + c_i\overline{Y}_{t-1} + d_i\Delta\overline{Y}_t + \varepsilon_{it}$, with $\overline{Y}_{t-1} = \sum_{i=1}^{N} Y_{it-1}$, $\Delta\overline{Y}_t = \sum_{i=1}^{N}\Delta Y_{it}$ and ε_{it} is the error term in the regression. For computational reasons, Pesaran (2007) advocates the use of a truncated version, $\overline{CADF}*$, where for positive constants K_1 and K_2 such that $Pr[-K_1 < CADF_i < K_2]$ is sufficiently large values of $CADF_i$ smaller than K_1 or larger than K_2 are replaced by the respective bound. Pesaran (2007) provides values for K_1 and K_2 as well as critical values for the test statistics obtained via stochastic simulation. The appropriate lag-length for $\overline{CADF}*$ is selected using the Akaike information criterion with a maximum equal to 4.

Our test results indicate that the common factors and the idiosyncratic components are $I(1)$ for all variables, as in 2. Therefore Pedroni approach is applied and the panel cointegration results are reported in Table 5.[14]

Reading across the rows it is evident the presence of cointegration between government spending, consumptions and relative price variables. Once evidence of a cointegrating relationship is found, we estimate the parameters of the Eq. (8) using the FM-CUP estimator proposed by Bai and Kao (2006):

$$\begin{aligned}\hat{\beta}_{CUP} &= \left[\sum_{i=1}^{n}\left(\sum_{i=1}^{T}\hat{y}_{i,t}^{+}\left(\hat{\beta}_{CUP}\right)\left(x_{i,t}-\tilde{x}_i\right)' - T\left(\hat{\lambda}_i'\left(\hat{\beta}_{CUP}\right)\hat{\Delta}_{F\varepsilon i}^{+}\left(\hat{\beta}_{CUP}\right) + \hat{\Delta}_{\mu\varepsilon i}^{+}\left(\hat{\beta}_{CUP}\right)\right)\right)\right] \\ &\times\left[\sum_{i=1}^{n}\sum_{t=1}^{T}\left(x_{i,t}-\overline{x}_i\right)\left(x_{i,t}-\overline{x}_i\right)'\right]^{-1}\end{aligned} \tag{16}$$

where $\hat{y}_{i,t}^{+} = y_{i,t} - \left(\hat{\lambda}_i'\hat{\Omega}_{F\varepsilon i} + \hat{\Omega}_{\mu\varepsilon i}\right)\hat{\Omega}_{\varepsilon i}^{-1}\Delta x_{i,t}$ indicates the transformation of the original dependent variable in order to correct for endogeneity, and $\hat{\lambda}_i$ the estimated factor loadings. The CUP-FM is constructed by estimating parameters, long-run covariances matrix (Ω) and factor loadings recursively. Thus $\hat{\beta}_{FM}$, $\hat{\Omega}$ and $\hat{\Lambda}_i$ are estimated repeatedly, until convergence is reached.[15] The Dynamic SUR estimator developed by (Mark et al., 2005) is also applied to provide more robust results.[16]

In Table 5 the parameter estimates and the elasticities of substitution are reported. The parameter estimates have the sign predicted by the theory. Hence the model yields plausible intertemporal (and intratemporal) elasticity of substitution. Specifically, $\nu/\alpha > 1/\alpha$, then the intratemporal elasticity of substitution dominates over the intertemporal elasticity of substitution indicating that government and private

[13] In this empirical analysis, we consider two regressors. However, the framework used by Gengenbach et al. (2006) leads to panel statistics for the null of no-cointegration that have the same distribution as panel unit root tests and hence are not affected by the number of regressors.

[14] This alternative is suggested by the nature of Bai and Kao (2006) CUP-FM estimator.

[15] In a Monte Carlo simulation, Bai and Kao (2006) show that this estimator has better small-sample properties than the two step-FM (2S-FM) and OLS estimator.

[16] We thank one of the referee for suggesting us the use of the Dynamic SUR estimator.

Table 5
Panel cointegration and estimation results.

Cointegration	Tests	Statistics
	Z_ρ	−6.762 (0.000)
	Z_t	−6.750 (0.000)
Estimation (CUP-FM)	Variables	Estimates [st. errors]
	$p_{i,t}$	−0.592 [0.143]
	$c_{i,t}$	1.324 [0.427]
	Elasticities	
	ν/α	0.755
	$1/\alpha$	0.447
Estimation (D-SUR)	Variables	Estimates [st. errors]
	$p_{i,t}$	−0.421 [0.156]
	$c_{i,t}$	1.123 [0.340]
	Elasticities	
	ν/α	0.890
	$1/\alpha$	0.375

Notes: Z_ρ and Z_t denote the panel coefficient ρ type and t-ratio tests. *p*-values are in brackets. For FMOLS estimators, the ICp2 information criteria are used to select the number of common factors. The maximum number of factors is fixed to 4. The standard errors are reported in parentheses. The CUP-FM *t*-statistics is well approximated by a standard N(0,1). D-SUR indicates the Dynamic SUR estimator proposed by Mark et al. (2005).

spending are Edgeworth-Pareto substitutes. Different results have been found in the literature. (Nieh and Ho, 2006) using a larger panel data set of countries found evidence of complementarity. However we believe that the selection of different countries may be responsible for such differences. The countries in the panel of (Nieh and Ho, 2006) are, for instance, less homogeneous than the EU countries.

4. Conclusions

In the European Union (EU) fiscal policy has become the main instrument available to national authorities as a stabilization policy. Consequently, issues related to fiscal policy have gained a growing interest, 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. 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 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. This paper empirically verifies the extent of direct substitution between government spending and private consumption in 15 EU countries using recently developed non-stationary panel methodologies that assume cross-section dependence. We consider a two-goods permanent-income model which allows us to estimate both the intratemporal and intertemporal elasticities of substitution between the two types of expenditure. In a preliminary time series analysis, the intratemporal elasticities of substitution (ν/α) of 15 EU member countries are found to be quite similar, but the intertemporal elasticities of substitution ($1/\alpha$) are rather diverse, implying that the substitutability between private and government consumption varies among the EU countries. When we assume cross-country dependence, we found that, on average, there is substantial substitutability between private and government consumption, implying that there will be direct crowding out of private consumption by government consumption. Although this relationship of substitutability between public and private consumption has important implications for assessing the overall effectiveness of fiscal policy, we need to be careful in advising policy maker since the empirical results are obtained using aggregate data.

References

Amano, R., Wirjanto, T., 1997. Intratemporal substitution and government spending. Review of Economics and Statistics 79 (4), 605–609.

Amano, R., Wirjanto, T., 1998. Government expenditures and the permanent-income model. Review of Economic Dynamics 1 (3), 719–730.

Aschauer, D.A., 1985. Fiscal policy and aggregate demand. American Economic Review 75 (1), 117–127.

Bai, J., Kao, C., 2006. On the estimation and inference of a panel cointegration model with cross-sectional dependence. In: Baltagi, B.H. (Ed.), Panel Data Econometrics: Theoretical Contributions And Empirical Applications. Elsevier Science & Technology, pp. 3–30. chapter 1.

Bai, J., Ng, S., 2004. A panic attack on unit roots and cointegration. Econometrica 72 (4), 1127–1177.

Campbell, J. Perron, P. 1991. Pitfalls and opportunities: What macroeconomists should know about unit roots, *NBER Technical Working Papers 0100*, NBER.

Choi, I., 2001. Unit root tests for panel data. Journal of International Money and Finance 20 (2), 249–272.

Djajic, S., 1987. Government spending and the optimal rates of consumption and capital accumulation. The Canadian Journal of Economics 20 (3), 544–554.

Elliott, G., Rothenberg, T., Stock, J., 1996. Efficient tests for an autoregressive unit root. Econometrica 64, 813–836.

Garber, P. M. King, R. G. 1983. Deep structural excavation? a critique of Euler equation methods, *National Bureau of Economic Research Technical Working Paper Series* No. 31.

Gengenbach, C., Palm, F.C., Urbain, J.P., 2006. Cointegration testing in panels with common factors. Oxford Bulletin of Economics and Statistics 68 (s1), 683–719.

Karras, G., 1994. Government spending and private consumption: some international evidence. Journal of Money, Credit & Banking 26 (1), 9–22.

Kwan, Y. K. 2006. The direct substitution between government and private consumption in east Asia, *National Bureau of Economic Research Working Paper Series* No. 12431.

Leybourne, S.J., McCabe, B., 1994. A consistent test for a unit root. Journal of Business and Economic Statistics 12 (2), 157–166.

Maddala, G., Wu, S., 1999. A comparative study of unit root tests with panel data and a new simple test. Oxford Bulletin of Economics and Statistics 1, 631–652.

Mark, N.C., Okagi, M., Sul, D., 2005. Dynamic seemingly unrelated cointegrating regression. Review of Economic Studies 72, 797–820.

Ni, S., 1995. An empirical analysis on the substitutability between private consumption and government purchases. Journal of Monetary Economics 36 (3), 593–605.

Nieh, C., Ho, T., 2006. Does the expansionary government spending crowd out the private consumption? Cointegration analysis in panel data. Quarterly Review of Economics and Finance 46 (1), 133–148.

OECD Statistical Compendium, 2007. Compendium Des Données Statistiques De L'ocde. DSI Data Service & Information, Dept. CD-ROM, 2007.

Ogaki, M., 1992. Engel's law and cointegration. Journal of Political Economy 100 (5), 1027–1046.

Ogaki, M., Park, J., 1997. A cointegration approach to estimating preference parameters. Journal of Econometrics 82 (1), 107–134.

Ogaki, M., Reinhart, C., 1998. Intertemporal substitution and durable goods: long-run data. Economics Letters 61, 85–90.

Okubo, M., 2003. Intratemporal substitution between private and government consumption: the case of Japan. Economics Letters 1, 75–81.

Pedroni, P., 2001. Purchasing power parity tests in cointegrated panels. Review of Economics and Statistics 83 (4), 727–731.

Pedroni, P., 2004. Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the ppp hypothesis. Econometric Theory 20 (3), 597–625.

Pesaran, M., 2007. A simple panel unit root test in the presence of cross section dependence. Journal of Applied Econometrics 22 (2), 265–312.

Pesaran, M. H. 2004. General diagnostic tests for cross section dependence in panels, *Cesifo Working Paper series 1240*, IFO.

Phillips, P.C.B., Hansen, B., 1990. Statistical inference in instrumental variables regression with I(1) process. Review of Economics Studies 57, 99–125.

Phillips, P.C.B., Ouliaris, S., 1990. Asymptotic properties of residual based tests for cointegration. Econometrica 58 (1), 165–193.

Prohl, S., Westerlund, J., 2009. Using panel data to test for fiscal sustainability within the European Union. FinanzArchiv: Public Finance Analysis 2, 246–269.

Redoano, M. 2003. Fiscal interactions among European countries, *Warwick Economic research papers*.

Van Dalen, H., 1999. Intertemporal substitution in public and private consumption — long-run evidence from the US and the UK. Economic Modelling 3, 355–370.