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BALANCING THE BUDGET THROUGH REVENUE OR SPENDING ADJUSTMENTS? THE CASE OF GREECE

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This paper examines the solvency of the Greek fiscal policy. Employing a cointegrated VAR as a benchmark, evidence of a long-run link between revenues and spending is presented, although intertemporal solvency is violated. Utilizing Granger-causality tests, a test for fiscal adjustment neutrality and Generalized Impulse Responses, this paper provides evidence in favor of the 'tax and spend' hypothesis for Greece. Additionally, the empirical evidence indicates that fiscal adjustment should take place through spending rather than revenue adjustment.

Keywords: Budget Balance, Government Revenue and Spending, Causality, Generalized Impulse Responses, Greece
JEL classification: C32, E60, H60

1. INTRODUCTION

The issue of curtailing budget deficits is one of the central themes of economic policy for many member countries of the European Union (EU). Two of the key requirements for European Monetary Union (EMU) membership are that budget deficit is a maximum of 3% of national income, and that the amount of government borrowing should not exceed 60% of national income, measured as real GDP. Hence correcting for fiscal imbalances is a necessary condition for EMU membership.

The issue of fiscal solvency of the government's financial policy - that is whether it satisfies its intertemporal budget constraint - and the effects that budget deficits have motivated several empirical and theoretical studies. Most of these studies have been carried out for the U.S. economy, delivering mixed results.

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This study assesses if Greek fiscal policy is solvent, by employing cointegration analysis. More specifically, I find evidence of cointegration between government spending and revenues, but I find that the Greek budget has not historically satisfied the intertemporal solvency condition. To this end, I try address the issue of how to achieve intertemporal government budget balance, i.e., through revenue increases or spending reductions.[1]

More specifically, given the evidence of cointegration, a vector equilibrium correction model (VEqCM) was formulated in order to explore the direction of causality. This is of interest from an economic point of view for two reasons. First, it allows one to acquire empirical knowledge about the long-run performance of the government. Second, based on the estimation results, it allows one to evaluate whether control over one variable, spending or taxation, can lead to control over the other variable or not. The approach taken here is the Johansen procedure, which is an appropriate framework for the analysis of causality, allowing for simultaneous investigation of both cointegration and Granger causality.[2]

Finally, in this econometric framework, I am able to address the question of relative efficiency of adjustments through taxes or expenditures. Essentially, I focus on the long-run effect of both means of adjustment and consider the issue of their 'neutrality' on long-run budget size and balance. The evidence presented show that the goal of balancing the budget could be achieved by reducing spending rather than increasing revenues.

The rest of the paper is organized as follows. Section 2 gives the conditions of intertemporal government solvency in the long run and discusses the main theoretical hypotheses between the variables of interest. Section 3 gives a brief review of the literature, while Section 4 contains a description of the data used in the analysis and the econometric framework employed. The empirical results are presented in Section 5. The final section contains some concluding remarks.

2. THE GOVERNMENT INTERTEMPORAL BUDGET CONSTRAINT

In every period the government must meet a budget constraint. Assuming that all government debt has one-period maturity, the government budget constraint can be written as:

[1] Bohn (1991) used an equilibrium correction model (EqCM) for the U.S. to describe how intertemporal budget imbalances were corrected in the past. Although this study uses similar methodology, the questions addressed are somewhat different. Specifically I evaluate whether adjusting either expenditures or revenues is 'neutral' for achieving equilibrium.

[2] The advantage of such an approach is that each variable is considered as potentially endogenous, being related to its own past values and the past values of other variables in the system.

$$G_t + (1 + i_t)B_t = R_t + B_{t+1}\,, \tag{1}$$

where G_t denotes government spending net of interest payments in period t, B_{t+1} are the funds raised by the government through bond sales in period t, i_t is the interest rate on the bonds sold in period $t-1$, and R_t are government's revenues in period t. The variables in Equation (1) can be nominal, real or deflated by population or GDP[3]. In what follows, I express these variables relative to GDP. dividing (1) by nominal GDP we have:[4]

$$\frac{G_t}{GDP_t} + \left(\frac{1 + i_t}{1 + \gamma_t}\right)\frac{B_t}{GDP_{t-1}} = \frac{R_t}{GDP_t} + \frac{B_{t+1}}{GDP_t}\,, \tag{1a}$$

where γ_t is the nominal GDP growth rate. Now letting lower case letters denote variable relative to GDP and let the net interest rate be $1 + \tilde{i}_t \equiv (1 + i_t)/(1 + \gamma_t)$, (1a) can be written as:

$$g_t + (1 + \tilde{i}_t)b_t = r_t + b_{t+1}\,. \tag{1b}$$

Solving (1b) forward yields the government's intertemporal budget constraint:

$$b_t = \sum\nolimits_{i=0}^{\infty} Q_{t,t+i}(r_{t+i} - g_{t+i}) + \lim_{l \to \infty} Q_{t,t+l}b_{t+l+1}\,, \tag{2}$$

where $Q_{t,t+i} = \prod\nolimits_{j=0}^{i}(1 + \tilde{i}_{t+j})^{-1}$. If the limit term in (2) equals zero then the current stock of outstanding debt, b_t, is just equal to the discounted value of future government surpluses. This condition is often called the 'No Ponzi Game' condition and implies that the government cannot retain its debt simply by issuing new debt perpetually.

If it is assumed that the net interest rate is stationary[5] with mean of $\tilde{i}$, then (1b) can be rewritten as,

[3] The interpretation of the interest rate in Equation (1) depends upon how the other variables are measured. See Hakkio and Rush (1991) fn. 2.

[4] See also Kremers (1989) for a discussion.

[5] Here I implicitly assume that the Fisher parity holds. In this case the net interest rate can be written as $(i_t - \Delta p_t) - \Delta y_t$, where Δp_t is the inflation rate and Δy_t is the growth rate of real GDP. Of course, as an anonymous referee of this journal pointed it out, the net interest rate can be non-stationary, but only if the Fisher parity condition fails. Similarly, if $(i_t - \Delta p_t)$ is stationary, with Δy_t being also stationary, then the net interest rate $\tilde{i}_t$ will also be stationary, as I have assumed.

$$g_t + \left(\tilde{i}_t - \tilde{i}\right)b_t + \left(1 + \tilde{i}\right)b_t = r_t + b_{t+1}\,, \tag{3}$$

which must hold every period. Define $e_t = g_t + \left(\tilde{i}_t - \tilde{i}\right)b_t$, i.e., government spending inclusive of interest payments. Substituting and solving (3) forward yields Equation (4):

$$b_t = \sum_{i=0}^{\infty} \rho^{i+1}\left(r_{t+i} - e_{t+i}\right) + \lim_{l\to\infty} \rho^{l+1} b_{t+l+1}\,, \tag{4}$$

where $\rho = \left(1 + \tilde{i}\right)^{-l}$. Using Equation (1b) and taking first differences of (3), Equation (4) can then be rewritten as,

$$s_t \equiv r_t - ge_t = \sum_{i=1}^{\infty} \rho^{i}\left(\Delta e_{t+i} - \Delta r_{t+i}\right) - \lim_{l\to\infty} \rho^{l+1} \Delta b_{t+l}\,, \tag{5}$$

where $ge_t \equiv g_t + \tilde{i}_t b_t$ denotes government expenditure, r_t government revenue and $s_t \equiv r_t - ge_t$ the budget surplus, all relative to GDP. Now if ge_t and r_t are integrated of order one $(I(I))$, so that their first differences are stationary, the right hand side of (5) is stationary. Imposing further the restriction that the growth rate of debt, Δb_t, is stationary implies that the limit term in (5) goes to zero.[6] Thus the restriction implied by the government's intertemporal budget constraint is that $r_t - ge_t$ must be stationary or that these two series are cointegrated with cointegrating vector of $[1, -1]$.[7] Furthermore, existence of a cointegrating relationship itself implies that ge_t and r_t share a common stochastic trend (Stock and Watson (1988)).

One observation should be made at this point. Notice, that if the source of nonstationarity is due to deterministic trends, say in the relative-to-output revenue and expenditure measures, the deterministic trend in the revenue-to-GDP ratio must be the same as that in the expenditure-to-GDP ratio (inclusive of interest payments) or the deficit will grow unboundedly as a proportion of GDP. Equation (5) then also implies a common deterministic trend.

[6] This is actually a stronger restriction than that which is necessary for the limit term to vanish as demonstrated by Quintos (1995). But relaxing this restriction implies eventual government default on its outstanding stock of debt. Given this outcome, if the growth rate of debt is not stationary, the government might be unable to market its debt. Since the growth rate of real output is stationary, if $\Delta b_t = \Delta\left(\left(B_t / P_t\right)/\left(GDP_t / P_t\right)\right)$ is nonstationary the debt-to-output ratio is growing. Furthermore, since taxes are, presumably, bounded as a percentage of real output, a growing debt-to-output ratio implies eventual default by the government.

[7] Hakkio and Rush (1991) and Quintos (1995) demonstrate that the cointegration vector $[1 - b]$, where $0 < b < 1$, is consistent with deficit sustainability since the limit term in (5) would still go to zero, but this would imply eventual government default on its debt, assuming no change in fiscal policy.

Another important issue to investigate is the existence and the direction of Granger-causality between the two variables of interest. In general, there are four hypotheses regarding the direction of causality between government spending and revenues that have been used in the literature: (a) spending precedes revenue; (b) revenue precedes spending; (c) revenue and spending are jointly determined; and (d) revenue and spending are independent of each other. There are several theoretical motivations behind each hypothesis.

1. *Spend and Tax.* Peacock and Wiseman (1979) and Barro (1979) argue that spending precedes revenue. More specifically, the tax smoothing hypothesis of Barro (1979) takes the path of government spending as given and taxes are adjusted to minimize distortions, while the budget is balanced intertemporally.

2. *Tax and Spend.* Friedman (1978) argues that a government adjusts spending to the level of revenue, and therefore higher taxes would increase public expenditures. In this case, imposing higher taxes to restrict the size of the government deficit would instead raise it.

3. *Spending Changes Simultaneously with Revenue.* This hypothesis is based on the equivalence of marginal cost and marginal revenue of the utility-maximizing suppliers and demanders of the public services. In this case the two aggregates mutually reinforce each other.

4. *Spending and Revenue Change Independently of each other.* According to Hoover and Sheffrin (1992), the level of spending and revenue can be set by a rule of thumb, reflecting the view of institutional separation of allocation and taxation functions of the government.

Which of these hypotheses is supported for Greece is an empirical question. Additionally, the issue of how it is more efficient to achieve intertemporal budget balance is important since it crucially hinges on which of the above assumptions hold. This issue will be further discussed in Section 4, where the econometric methodology is laid out.

3. PREVIOUS STUDIES

There are many studies that have been performed on the government's intertemporal budget and the causal relationship between revenue and government spending. In a seminal article, Hamilton and Flavin (1986) demonstrate that the federal government cannot run a permanent deficit exclusive of interest payments on the debt, but may have a constant deficit when interest payments are included. The key result is that the real rate of interest at which the government borrows must be greater than the growth rate of real debt so that the discounted present value of future government debt goes to zero. Furthermore, they claim that if the government budget is to be balanced in present value terms, the surplus (deficit) inclusive of interest payments should be stationary. Hamilton and Flavin present evidence that this condition is satisfied using annual data from 1960 to 1984. But Kremers (1988) demonstrates that this result is not robust to lag specification in the test. Trehan and Walsh (1988) generalize the Hamilton and Flavin result to show that government expenditures, inclusive of interest payments, and revenues should be cointegrated with a cointegration vector equal to $[1,-1]'$. They present evidence that supports this restriction. Hakkio and Rush (1991) point out that the restriction that the cointegrating vector between government expenditures, inclusive of interest payments, and revenues equal $[1,-\beta_{ge}]'$, such that $\beta_{ge}=1$, is a sufficient condition for intertemporal budget balance but not, strictly speaking, a necessary condition. Quintos (1995) expands on Hakkio and Rush (1991) by stating the necessary and sufficient conditions for deficit sustainability as $0<\beta_{ge}\leq 1$. Haug (1995) presents similar results, using a slightly different methodology.

On the other hand, Owoye (1995) examines the causal relationship between taxes and spending in the G7 countries by applying equilibrium-correction models. According to this study, bi-directional causality exists between taxes and spending in five G7 countries, but unidirectional causality from revenues to spending exists only in the case of Italy and Japan. Ram (1988) finds that taxes Granger-cause spending at the federal level for the U.S., but spending Granger-causes taxes at the state and local level. Koren and Stiansy (1998) provide mixed results in a panel of nine industrialized countries. They find that in that causality runs in either direction depending on the country under study. Hondroyiannis and Papapetrou (1996) examine the Greek public finances; using a VEqCM, they find that taxes Granger-cause government spending, therefore that the observed behavior of the Greek government is consistent with the *tax and spend hypothesis*. Finally, in a related study, Kollias and Makrydakis (2000), find that there is bi-directional causality between revenues and taxes for Greece and Ireland, whereas the *tax and spend* hypothesis is supported for Spain and no direction of causality is established for Portugal.

4. DATA AND ECONOMETRIC METHODOLOGY

4.1 Data

The data used in this study are quarterly from 1970:1 to 1997:1. Unfortunately, to the best of my knowledge, a longer time span of data on the Greek government budget is not available. The government expenditures inclusive of interest payments and government revenues were sampled from the *International Financial Statistics* (IFS)

Database. The nominal GDP series were obtained from the OECD *Business Sector Data Base* (BSDB). Government revenues and expenditures were seasonally adjusted using the Census X-11 (multiplicative) method.

I use nominal variables relative to nominal GDP, since it is difficult to identify the appropriate deflator for expenditures and revenues, and the use of an inappropriate deflator could contaminate my results. Additionally, the employed theoretical framework can easily accommodate variables relative to GDP without any additional requirements, as was explained above.

4.2 Econometric Methodology

In order to test the hypotheses discussed in Section 2, Johansen's Full Information Maximum Likelihood (FIML) will be employed (Johansen (1995)). Johansen's procedure starts with the definition of an *n*-dimensional vector of variables x_t. The Vector Autoregressive (VAR) representation of the unrestricted system with k lags and Gaussian errors u_t is:

$$x_t = \sum_{i=1}^{k} A_i x_{t-i} + \mu + u_t , \tag{6}$$

where $u_t \rightarrow N(0,\Omega)$, x_t is a $(n \times 1)$ vector and each of the A_i is a $(n \times n)$ matrix of parameters. Model (6) can be reformulated into a Vector Equilibrium Correction (VEqCM) form:

$$\Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} + \mu + u_t , \tag{7}$$

where $\Pi = \sum_{i=1}^{k} A_i - I_n$, $\Gamma_i = -\sum_{j=i+1}^{k} A_j$ and for later reference define also $\Gamma = I_n - \sum_{i=1}^{k-1} \Gamma_i$.

The hypothesis that x_t is $I(1)$ is formulated as the reduced rank hypothesis of the matrix Π (see Johansen (1995)). Here one assumes that the matrix Π has rank $r < n$. In this case Π can be decomposed as the product of two matrices $\alpha\beta'$ where α, β are each $n \times r$ and have full rank $r < n$

$$\Pi = \alpha\beta' , \tag{8}$$

Furthermore, the full rank of $\alpha_\perp' \Gamma \beta_\perp$, is required, where $\alpha_\perp$ and $\beta_\perp$ are $n \times (n - r)$ matrices orthogonal to α and β respectively. Following this parameterization, there are r linearly independent stationary relations given by the cointegrating vectors β; the matrix α gives the speed of adjustment of the endogenous variables to their steady state values (the cointegrating relations), while there are also $n-r$ linearly independent non-stationary relations. These last relations define the common stochastic trends of the system. Under these restrictions, the solution of x_t as a function of the disturbances u_t, the initial conditions x_0, and the deterministic variables μ is given by

$$x_t = C(1)\sum_{t=1}^{T}(u_i + \mu) + C^*(L)(u_t + \mu) + \Phi\ , \tag{9}$$

where $C(1) = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}\alpha'_\perp$, $C\ (L)$ is a polynomial in the lag operator, and Φ is a function of initial conditions, such that $\beta'\ \Phi = 0$.

Regressing Δx_t and x_{t-1} on $\mu, \Delta x_{t-1}, ..., \Delta x_{t-k+1}$ gives the residuals R_{0t} and R_{1t} from which the residual product-moment-matrices $S_{ij} = T^{-1}\sum_{t=1}^{T} R_{it}R'_{\text{jt}}\ ; i, j = 0{,}1,$ are constructed. Then β is estimated as the eigenvectors associated with the *r* largest eigenvalues $1 > \lambda_1 > ... > \lambda_r > 0$ found as the solution to the eigenvalue problem: $\left|\lambda S_{11} - S_{10}S_{00}^{-1}S_{01}\right| = 0$. Johansen has developed a sequence of Likelihood Ratio (*LR*) tests for the hypothesis on the number of the cointegrating vectors (or equivalently the rank of Π); the so-called maximum eigenvalue test based on the statistic $Q(r|r+1) = -T\ln(1-\lambda_{r+1})$ and the trace test based on the statistic: $Q(r|n) = -T\sum_{i=r+1}^{n}\ln(1-\lambda_i)$. Doornik (1998) has proposed a way for calculating the asymptotic *p-values* for the above tests, while asymptotic critical values are given in Johansen (1995) *inter alia*.

4.3 Testing the Neutrality of Fiscal Adjustments

In a recent paper Garcia and Henin (1999) have proposed a way to evaluate the neutrality of adjusting fiscal imbalances by revenue or spending adjustments. The evaluation utilizes the VEqCM representation of the system $\Gamma(L)\Delta x_t = \mu + \alpha\beta' x_{t-1} + u_t$. The Wold representation of the system is given by:

$$\Delta x_t = C(L)u_t + \kappa = \left(C(1) + (1-L)\,C^*(L)\right)u_t + \kappa\,, \tag{10}$$

where $C(1)$ was given above in (9). Taking into account the fact that the matrix $C(1)$ has reduced rank, assuming that $\beta' = [1, -\beta_{ge}]$ and using the closed form solution in (9) it easy to see that

$$C(1) = K\begin{pmatrix} \beta_{ge}\alpha_{ge} & -\beta_{ge}\alpha_r \\ \alpha_{ge} & -\alpha_r \end{pmatrix} \text{ with } K = (\alpha'_\perp \Gamma \beta_\perp)^{-1}. \tag{11}$$

The long-run response of the government surplus $s_t \equiv r_t - ge_t$ to a revenue innovation is

$$C_{s,r}(1) = C_{r,r}(1) - C_{ge,r}(1) = K(\beta_{ge} - 1)\alpha_{ge}\,, \tag{12}$$

and the response to an expenditure innovation is

$$C_{s,ge}(1) = C_{r,ge}(1) - C_{ge,ge}(1) = K(1 - \beta_{ge})\alpha_r\,, \tag{13}$$

Then for the choice of a fiscal adjustment strategy to be neutral, the sum of both responses must be zero: a 1% increase in revenues should have the same long-run effect on the surplus as expenditure cuts of the same amount:

$$C_{s,r}(1) + C_{s,ge}(1) = K(\beta_{ge} - 1)(\alpha_{ge} - \alpha_r) = 0\,, \tag{14}$$

From Equation (14) it is clear that neutrality prevails if either (i) $\beta_{ge} = 1$ or $\beta' = [1,-1]$, since in this case no shock has persistent effects on the budget surplus, or if (ii) $\alpha_r = \alpha_{ge}$. Although both conditions are sufficient for neutrality they are associated with different mechanisms. In the first case, the stationarity of the budget balance implies its long-run invariance with respect to any perturbation. In the second case, the equilibrium correction term $\beta' x_{t-1}$ impacts the same way on revenues and expenditures, with no resulting effect on the surplus.

4.4 Generalized Impulse Responses

A standard way of judging the interaction of variables in a cointegrated system is to conduct impulse response analysis,[8] as has been advocated by Lütkepohl and Reimers (1992). Usually, in such an analysis the cointegrated VAR is subject to an orthogonalized shock in one variable and the response of the system is examined. Of course, this approach suffers, due to the dependence of the results on identifying assumptions employed to obtain the orthogonalized shocks.

Recent results of Koop *et al.* (1996) and Pesaran and Shin (1998) though, have re-examined the concept of impulse response analysis, aiming to remove this shortcoming. Instead of using orthogonalized impulse responses, one may use generalized impulse responses (GIR) that are based on a 'typical' shock to the system. The average response of the system to this typical shock is compared to the average baseline model where the shock is absent. Rather than examining the effect of a pure orthogonalized shock to, say, revenues, GIR analysis considers a typical *historical* innovation, which embodies information on the contemporaneous correlations between the innovations.

[8] See also Sims (1980) for an original contribution on the use of impulse response analysis.

The argument about GIR may be explained as follows. Let the Vector Moving Average (VMA) representation of the *n*-variable cointegrated VAR model be given by

$$\Delta x_t = \kappa_0 + \kappa_1 t + \sum_{i=0}^{+\infty} C_i u_{t-i} , \tag{15}$$

where κ_o is a vector of constants, κ_1 are the coefficients of the deterministic trend, and u_t is a vector of unobserved "shocks", with variance matrix Ω and let ω_{ij} be a typical element of Ω. Then it holds that

$$E\left(u_t \middle| u_{jt} = u_{jt}^0\right) = \omega_{jj}^{-1} u_{jt}^0 \Omega e_j , \tag{16}$$

where e_j is a $(n \times 1)$ selection vector with element j equal to unity and zeroes elsewhere. Then the GIR of the effect of a "unit" shock to the *j-th* disturbance term at time *t* on Δx_{t+h} is

$$GIR_{\Delta x_{t+h}}(h, j) = \frac{C_h \Omega e_j}{\sqrt{\omega_{jj}}}, h = 0,1,2,..., \tag{17}$$

and the GIR of x_{t+h} following a shock to the *j-th* variable is

$$GIR_{x_{t+h}}(h, j) = \frac{Q_h \Omega e_j}{\sqrt{\omega_{jj}}}, \; h = 0,1,2,..., \tag{18}$$

where $Q_h = \sum_{i=0}^{h} C_i$ and the GIRs are measured h periods after the shock has occurred.[9] It similarly follows that

$$GIR_{\beta' x_{t+h}}(h, j) = \frac{\beta' Q_h \Omega e_j}{\sqrt{\omega_{jj}}}, \; h = 0,1,2,... \tag{19}$$

As shown in Pesaran and Shin (1998), the GIR will be numerically equivalent to the standard impulse response function based on Cholesky decompositions if Ω is diagonal.

[9] In general, Pesaran and Shin (1998) show that one can interpret generalized impulse responses for a stationary vector process y_t as $GIR_{y_{t+h}}(\delta, \Im_{t-1}) = E\left(y_{t+h} \middle| u_t = \delta, \Im_{t-1}\right) - E\left(y_{t+h} \middle| \Im_{t-1}\right)$. They also explain that in a linear system, the impulse responses will be invariant to history (the information set on which conditioning is made), and so the GIR will depend only on the composition of the shocks as defined by δ .

5. EMPIRICAL RESULTS

5.1 Unit Root and Stationarity Tests

In order to see if there exists any long-run relationship between the variables of interest, I first tested for the order of integration of each variable. Since its is by now well known that the augmented Dickey-Fuller (ADF) tests have very low power, I also relied on ADF-type tests proposed by Elliott *et al*. (1996), denoted DF-GLS, which are based on prior detrending of the data using GLS and then estimating the ADF-type regression. I also used stationarity tests proposed by Kwiatkowski *et al.* (1992), denoted KPSS, where stationarity is taken as the null hypothesis, against the alternative of a unit root. These three tests make a good combination, as the null in the first two is the alternative of the latter. The results are reported in Table 1.

Table 1. Unit Root and Stationarity Tests

Panel A: Unit Root and Stationarity Tests for the Levels				
	r_t	ge_t	s_t	CV 5%
ADF	-0.356 (3)	0.170 (12)	-0.583 (12)	-2.889
DF-GLS	0.803 (3)	0.376 (10)	0.171 (12)	-1.943
KPSS	1.130 (9)	1.258 (8)	1.158 (8)	0.463
Panel B: Unit Root and Stationarity Tests for the First Differences				
	Δr_t	Δge_t	Δs_t	CV 5%
ADF	-17.132 (0)	-16.358 (0)	-16.075 (0)	-2.889
DF-GLS	-2.219 (10)	-15.672 (0)	-2.781 (5)	-1.943
KPSS	0.053 (8)	0.211 (28)	0.216 (33)	0.463

Note: The numbers in parentheses are the lags used in the ADF-type tests chosen by the modified AIC (Ng and Perron (2001)) or the bandwidth parameter for the KPSS tests.

All three tests indicated that the revenue and expenditure series are consistent with the hypothesis of a unit root. Additionally, I have examined the stationarity properties of the budget surplus, $s_t = r_t - ge_t$, which was also found to be consistent with a unit root. This constitutes *prima facie* evidence that the Greek government has not followed a policy that is consistent with long-run solvency, and that would render the budget surplus stationary. This has the further implication that shocks will have a permanent effect on the budget surplus. This issue is further explored below.

5.2 Testing the Intertemporal Restriction

In order to test whether the series constitute a cointegrated system and evaluate the intertemporal restriction more rigorously in a multivariate framework, the Johansen procedure was employed. The VEqCM model was estimated by using four lags, including an unrestricted drift term. The choice of the lag-length was based on the Akaike information criterion (Akaike (1969)) and on tests for no autocorrelation present in the residuals.[10] The results are summarized in Table 2.

Table 2. Cointegration Tests System $x_t = (r_t \quad ge_t)'$

Panel A: Trace and Maximum Eigenvalue Tests							
Trace Test: $Q(r/n)$				**Maximum Eigenvalue Test:** $Q(r/r+1)$			
H_0	H_1	$Q(r/n)$	*p-value*	H_1	H_1	$Q(r/r+1)$	*p-value*
$r=0$	$r \geq 1$	17.44	[0.023]	$r=0$	$r \geq 1$	17.43	[0.013]
$r \leq 1$	$r=2$	0.01	[0.906]	$r \leq 1$	$r=2$	0.01	[0.906]

Panel B: Multivariate Autocorrelation Diagnostics	
Test Statistic	*p-value*
Autocorrelation LM(1)	0.023
Autocorrelation LM(4)	0.140

Panel C: Cointegrating Relation	
$\beta' x_t =$	$r_t - 0.471\, ge_t$
$(t-stat.)$	(-12.779)

Notes: The estimation was based on a VEqCM of order four (VAR(5)), with the number of lags chosen by the Akaike information criterion (AIC). The model included a constant term unrestricted. The Autocorrelation $\text{LM}(i) i = 1,4$ is a test for i-*th* order autocorrelation distributed as a $\chi^2(4)$. The asymptotic *p-values* for the cointegration test are calculated using the method of Doornik (1998).

Both the maximum eigenvalue and the trace statistic suggest that there is exactly one cointegrating vector. More specifically, the null of no cointegration is rejected at the 5% significance level (the associated *p-values* for the trace and maximum eigenvalue statistics are 0.023 and 0.013 respectively), whereas the null that the rank is one is not rejected by both tests. This provides evidence that government revenues and expenditures are tied together in the long run, sharing a common stochastic trend. More specifically, there is evidence of a unique statistical equilibrium, that works as an

[10] The assumption of no autocorrelation is the most crucial one for the Johansen (1995) Maximum Likelihood procedure. Other deviations from whiteness do not seem to severely influence cointegration inference (Gonzalo (1994)).

'attractor' for the variables.[11] The estimated long-run relationship is given in panel C of Table 2 (with *t*-statistic in parenthesis).

The estimated cointegration vector normalized on the revenues variable is $[1, -0.471]$. This long-run relationship allows one to evaluate whether the intertemporal restriction holds or not. A likelihood ratio (LR) test of the restriction that the cointegration vector is $[1, -1]$ yields a statistic of 14.39, which is distributed as a $\chi^2(1)$ variate. Thus, the restriction implied by Equation (5), is violated. I will return to this issue below, where evidence on the neutrality hypothesis is presented and this hypothesis is examined in more detail.

Another important issue, along with the existence of cointegration between the variables in the system, is the issue of stability of the long-run relationships through time, as well as the stability of the estimated coefficients of such relationships.[12] Hansen and Johansen (1999) have suggested methods for the evaluation of parameter constancy in cointegrated VAR models, utilizing estimates obtained from the FIML estimation procedure. The first test deals with the hypothesis of constancy of the cointegration space for a given cointegration rank. Hansen and Johansen (1999) have proposed a likelihood ratio test that is constructed by comparing the likelihood from each recursive sub-sample to the likelihood function calculated under the restriction that the cointegrating vectors estimated from the full sample fall within the space spanned by the estimated vectors of each individual sample. The test statistic is χ^2 distributed with $(n-r)r$ degrees of freedom (see also Hansen and Juselius (1995)). The second test examines the constancy of each individual element of the identified cointegrating vectors, β. Additionally, one can also exploit the fact that there is a unique relationship between the eigenvalues and the cointegrating vectors.[13] Therefore, when the cointegrating vectors have undergone a structural shift this will be reflected in the estimated eigenvalues. Hansen and Johansen (1999) have derived the asymptotic distribution as well as the asymptotic variance of the estimated eigenvalues. Finally, Hansen and Johansen (1999) have also proposed two versions of a *LM* test, the *Nyblom* $\sup Q_T^{(t)}$ and $meanQ_T^{(t)}$ tests, to examine the constancy of the cointegrating space. These last two tests have non-standard distributions and have to be simulated.

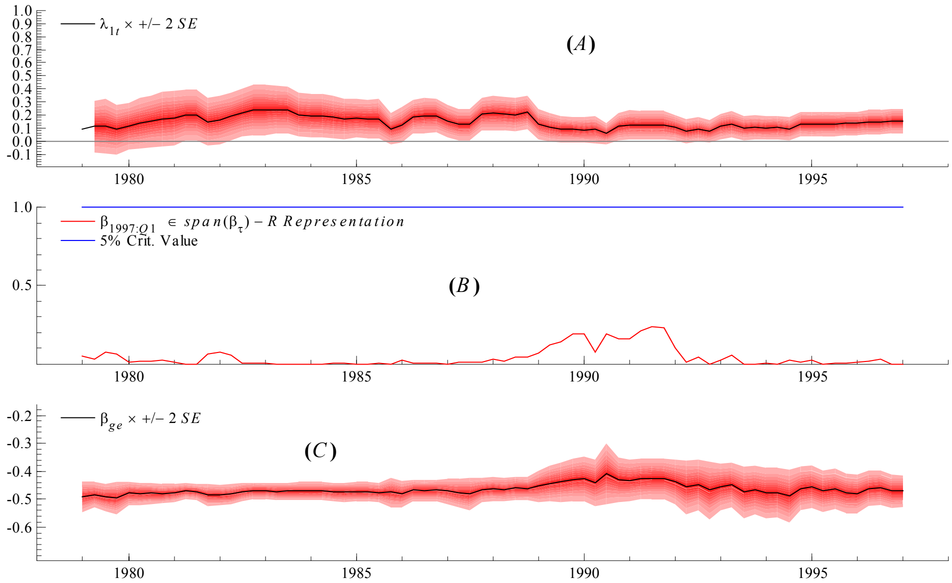
The graphical output of these tests is reported in Figures 1 and 2. In both figures, the reported results are based on recursive estimation of the model starting on the first quarter of 1979, due to the small sample available. First, Panels A and C of figure 1 show evidence that the recursive estimate of β_{ge} is stable, since both the recursive

[11] See also Johansen (1995) for a geometric interpretation of cointegration and the 'attractor' set.

[12] I am thankful to an anonymous referee of this journal for raising this issue on an earlier draft of the paper.

[13] Notice that $\hat{\alpha} = S_{01}\hat{\beta}$, hence $\hat{\alpha} S_{00}^{-1}\hat{\alpha} = \hat{\beta} S_{10} S_{00}^{-1} S_{01}\hat{\beta} = diag(\hat{\lambda}_1, ..., \hat{\lambda}_r)$ (see Johansen (1995), Chapter 6 for a detailed analysis with all the derivations of the formula).

coefficient estimate and the time path of the recursively estimated eigenvalue, are virtually flat. Second, Panel B of figure 1 shows an asymptotically χ^2 distributed test, where the null is that of a constant cointegrating space (Hansen and Juselius (1995)).



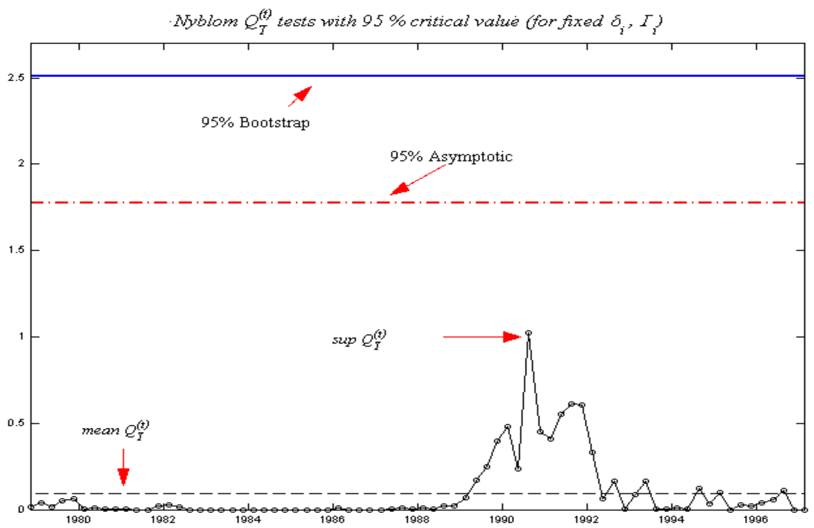
Note: Panel A reports the recursively-estimated non-zero eigenvalue, Panel B the test for stability of β ($\chi^2(2)$ -distributed) and Panel C the recursive estimates of β_{ge} .

Figure 1. Cointegrating Space Parameter Stability Tests

Again, the test statistic provides evidence in favor of the cointegrating space constancy null. Finally, Figure 2 displays the *Nyblom* tests proposed in Hansen and Johansen (1999), where it is also evident that there is no significant parameter instability associated with the cointegrating space.[14, 15]

[14] The 5% asymptotic critical value for the $\sup Q_T^{(t)}$ test is 1.776 and the corresponding bootstrapped critical value is 2.509. These critical values have been obtained by Monte Carlo simulation (10.000 replications). In addition, I have also found that $\sup Q_T^{(t)} = 1.02$ and $mean Q_T^{(t)} = 0.09$ with associated asymptotic *p-values* of [0.24] and [0.60] (and bootstrapped *p-values* of [0.21] and [0.58]) respectively.

[15] In addition to these parameter stability tests, I have experimented using dummies in the VEqCM model. More specifically, I have utilized two dummies for to account for the two political parties in power from 1974 onwards (New Democracy and PASOK). *LR* and *Wald* tests of the exclusion restrictions ($\chi^2(4)$ distributed)



Notes: The figure plots the *Nyblom* test $Q_T^{(t)}$ for the stability of (unrestricted) β along with the asymptotic (1.776) and bootstrapped (2.509) critical values (for the $\sup Q_T^{(t)}$). The critical values were obtained from 10,000 replications of the empirical model.

Figure 2. Nyblom Tests of Stability of the Cointegrating Space

5.3 Causality

Assessment of the direction of the causal nexus is an empirical issue. By causality, I mean causality in Granger's (1969) sense. That is I would like to examine whether one variable precedes the other, or if they are contemporaneous. In this case the question as to whether r_t causes ge_t amounts to testing how much of the current value of the second variable can be explained by past values of the first variable. So ge_t is said to be Granger-caused by r_t, if the coefficients of lagged r_t are statistically significant in the regression of ge_t on lagged values of all the variables in the information set. A test of Granger-causality can be empirically carried out by means of the VAR model (in

were 2.017 and 4.715 (*p-values* [0.733] and [0.318]) respectively, indicating no significant parameter change due to different political parties in power. I have also experimented using a dummy for the Maastricht treaty, and again both the *LR* and the *Wald* tests ($\chi^2(2)$ distributed) were 0.028 and 0.038 (*p-values* [0.986] and [0.981]) respectively, thereby also reflecting no significant parameter change.

levels) if the variables are stationary. However, if this is not the case, the test statistics will converge to non-standard distributions, and the empirical results can be misleading (Sims *et al.* (1990)). Instead, Granger-causality can be tested using a VEqCM as:

$$\Delta r_t = \mu_r + \sum_{i=1}^{k-1}\Gamma_{i,r}^{r}\Delta r_{t-i} + \sum_{i=1}^{k-1}\Gamma_{i,ge}^{r}\Delta ge_{t-i} + \alpha_r \beta' x_{t-1} + u_t^r \ , \tag{20}$$

$$\Delta ge_t = \mu_{ge} + \sum_{i=1}^{k-1}\Gamma_{i,r}^{ge}\Delta r_{t-i} + \sum_{i=1}^{k-1}\Gamma_{i,ge}^{ge}\Delta ge_{t-i} + \alpha_{ge} \beta' x_{t-1} + u_t^{ge} \ , \tag{21}$$

where u_t^r and u_t^{ge} are error terms, which are assumed to be white noise. According to Granger (1988), the presence of cointegration implies Granger-causality in at least one direction between the variables involved. A further distinction is that of *long-run* vs. *short-run* causality. If the values of $\Gamma_{i,r}^{ge}\left(\Gamma_{i,ge}^{r}\right)$ are jointly zero for all *i*, then the hypothesis that $r_t(ge_t)$ does not Granger-cause $ge_t(r_t)$ in the short run cannot be rejected, while if the value of $\alpha_{ge}\left(\alpha_r\right)$ is zero the hypothesis that $r_t(ge_t)$ does not Granger-cause $ge_t(r_t)$ in the long run cannot be rejected. Similarly, the joint hypotheses that $\Gamma_{i,r}^{ge}$ and α_{ge} $\left(\Gamma_{i,ge}^{r} \text{ and } \alpha_{\mathrm{r}}\right)$ are zero, implies Granger non-causality from $r_t(ge_t)$ to $ge_t(r_t)$.

Before examining the direction of Granger causality, I have estimated the unrestricted VEqCM (with the cointegrating vector unrestricted). The results along with some diagnostic tests are reported in Table 3.

The results show that both equations are well specified in terms of autocorrelation - although normality is ascent - with quite some predictive ability for each equation (relative to a random walk with drift benchmark). It is important also to point out that both variables seem to equilibrium correct significantly towards the underlying long-run relationship, while both adjustment coefficients have the expected sign.

The results from the Granger causality test are summarized in Table 4.

From Table 4, it is evident that there is long-run causality running in both directions, with the long-run non-causality from ge_t to r_t being only marginally rejected. On the other hand, expenditure does not Granger-cause revenue in the short-run, while the joint hypothesis of Granger non-causality from ge_t to r_t finds strong empirical support. Based on these findings, and recalling our discussion in Section 2, I can conclude that the Greek fiscal authorities have been behaving according to the *tax and spend hypothesis*, since revenues precede spending. This also implies that the Greek fiscal authorities should probably follow a fiscal adjustment, by reducing spending rather than increasing revenues. This is further explored below.

Table 3. VEqCM Estimates

Dependent Variable	Equation	
	Δr_t	Δge_t
Δr_{t-1}	-0.504	-1.826
$(t-stat)$	(-3.858)	(-3.098)
Δr_{t-2}	-0.306	-1.001
$(t-stat)$	(-2.265)	(-1.640)
Δr_{t-3}	-0.200	-1.047
$(t-stat)$	(-1.579)	(-1.828)
Δr_{t-4}	-0.036	-1.000
$(t-stat)$	(-0.343)	(-2.124)
Δge_{t-1}	-0.099	-0.303
$(t-stat)$	(-2.153)	(-1.450)
Δge_{t-2}	-0.085	-0.592
$(t-stat)$	(-1.980)	(-3.068)
Δge_{t-3}	-0.082	-0.280
$(t-stat)$	(-2.322)	(-1.751)
Δge_{t-4}	-0.025	-0.283
$(t-stat)$	(-0.983)	(-2.446)
$\beta' x_{t-1}$	-0.205	1.558
$(t-stat)$	(-1.892)	(3.127)
δ	0.004	-0.025
$(t-stat)$	(2.240)	(-2.824)
R^2	0.372	0.645
Adjusted R^2	0.312	0.611
Autocorrelation LM(4)	4.953 [0.292]	13.427 [0.010]
ARCH LM (4)	2.741 [0.602]	4.964 [0.291]
NORM $\chi^2(2)$	12.432 [0.002]	25.394 [0.000]

Notes: Δ stands for the first difference operator. The Autocorrelation *LM*(4) and ARCH *LM*(4) are distributed as $\chi^2(4)$.

Table 4. Granger Causality Tests

H_0	α_r / α_{ge}	$\Gamma_{i,ge}{}^{r} / \Gamma_{i,r}{}^{ge}$	$\alpha_r \cup \Gamma_{i,ge}{}^{r} / \alpha_{ge} \cup \Gamma_{i,r}{}^{ge}$
$ge_t \not\succ r_t$	$\chi^2(1) = 3.58$ [0.06]	$\chi^2(4) = 7.09$ [0.13]	$\chi^2(5) = 7.35$ [0.19]
$r_t \not\succ ge_t$	$\chi^2(1) = 10.16$ [0.00]	$\chi^2(4) = 12.11$ [0.02]	$\chi^2(5) = 15.04$ [0.01]
$ge_t \not\succ r_t \cap r_t \not\succ ge_t$	$\chi^2(2) = 17.15$ [0.00]	$\chi^2(8) = 21.09$ [0.01]	$\chi^2(10) = 25.89$ [0.00]

Notes: The symbol $\not\succ$ denotes Granger non-causality. The first hypothesis is that spending does not Granger-cause revenues, the second that revenue does not Granger-cause spending, and the last that spending does not Granger-cause and is not Granger-caused by revenues.

5.4 Assessing the Neutrality of Adjustments

In this subsection, I examine whether fiscal adjustments by revenue increase or spending decrease are neutral for balancing the budget. In addition to testing the two hypotheses outlined in Section 4, Garcia and Henin (1999) also suggest that one can evaluate the overall probability of neutrality as:[16] $\Pr(\beta_{ge} = 1) + \Pr(\alpha_r = \alpha_{ge}) - \Pr(\beta_{ge} = 1 \cap \alpha_r = \alpha_{ge})$. The results of these tests are presented in Table 5.

Table 5. Tests on the Cointegrating Vector and the Adjustment Coefficients

VAR(5)	Unrestricted	$H_0 : \beta = (1 \quad \text{-}1)'$	$H_0 : \alpha_r = \alpha_{ge}$	$H_0 : \beta = (1 \quad \text{-}1)' \cup \alpha_r = \alpha_{ge}$
Coint. Vector β (*t*-statistics)	$(1 \quad \text{-}0.471)$ $(\cdots, \text{-}12.779)$	$(1 \quad \text{-}1)'$	$(1 \quad \text{-}0.479)$ $(\ldots, \text{-}5.826)$	$(1 \quad \text{-}1)'$
Adjustment Coeff. Revenues α_r (*t*-statistic)	-0.204 (-1.89)	-0.026 (-0.893)	-0.201 (-1.876)	-0.023 (-0.806)
Adjustment Coeff. Expenditures α_{ge} (*t*-statistic)	1.556 (3.19)	0.169 (1.254)	-0.201 (-1.876)	-0.023 (-0.806)
LR $Q(v) \sim \chi^2(v)$		*Q*(1) = 14.39	*Q*(1) = 14.08	*Q*(2) = 16.73
p-value		[0.00]	[0.00]	[0.00]

[16] Formally speaking, the joint restrictions on β and α_r, α_{ge} is problematic, since for the VAR model with just one lag it implies that $\beta'_{\perp} = [1,1], \alpha'_{\perp} = [\alpha_r, -\alpha_r], \Gamma = I_n$, hence $|\alpha'_{\perp} \Gamma \beta_{\perp}| = |\alpha'_{\perp} \beta_{\perp}| = 0$ and process x_t becomes *I(2)* (see Johansen (1995, Ch.9)).

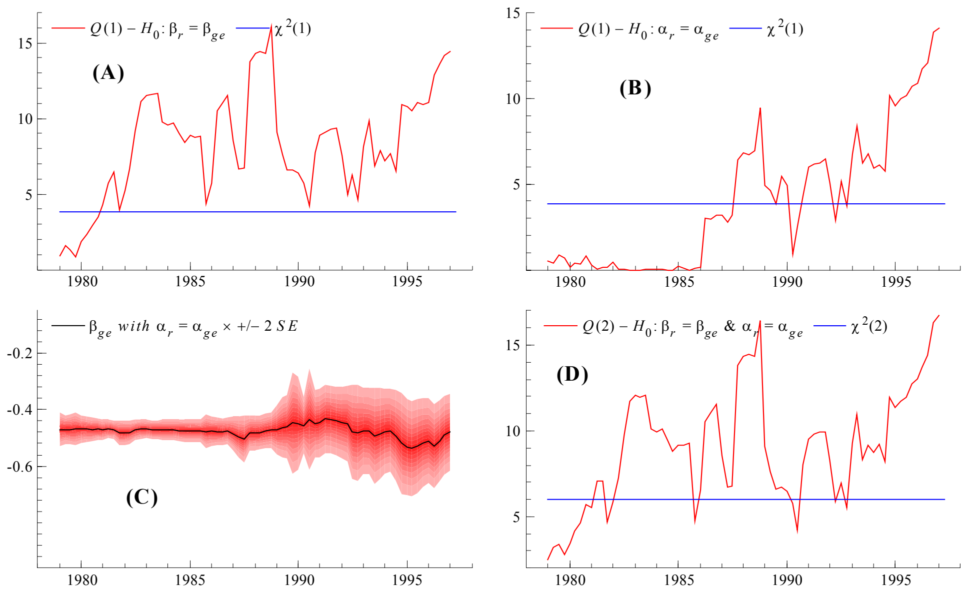
The first column of Table 5 contains the unrestricted estimates of the parameters of interest. The second column presents the parameters of interest under the restriction that the budget surplus is stationary, which delivers a *LR* test of 14.39, hence rejecting the null. Similarly, the third column presents the parameter estimates under the restriction of equality of the speed of adjustment of revenue and expenditure. The results indicate that this hypothesis is also strongly rejected, delivering a *LR* test of 14.08. Finally, the last column of presents a test of the joint hypothesis of balanced budget and equal adjustment, which is also strongly rejected. Overall my results indicate that the neutrality hypothesis does not hold for Greece.[17]

The results just presented, are based on the full sample period. One could of course imagine that some of these hypotheses hold in some sub-periods while they fail for others.[18] In order to entertain this possibility, I have recursively calculated all the above *LR* tests for a period starting from the first quarter of 1979. The graphical output is reported in Figure 3. Before proceeding with the discussion of the findings, a word of caution is necessary regarding the interpretation of the findings. These tests are asymptotically distributed as χ^2 variates. This holds for relatively large samples and under the assumption of Gaussian residuals, which, as we saw from Table 3, seems to be violated. Hence the results that follow should be interpreted as suggestive rather than definite, given the absence of residual normality.

Having put the analysis in perspective, notice that the restriction of budget balance seems to hold for the period prior to 1981, i.e., before the socialist government of PASOK came to power (Panel A, Figure 3). Similarly, the restriction of equal speeds of adjustment seems to hold for the period prior to 1986, and for some quarters in the early 1990s (Panel B, Figure 3). In addition, the estimate of the cointegrating parameter β_{ge} (Panel C, Figure 3) - under the restriction of equal speeds of adjustment - seems to be stable, albeit the associated confidence bands seem to be wider for the whole 1990s. Finally, as far as the joint restriction of balanced budget and equal speed of adjustment is concerned, it can clearly be seen from Panel D of Figure 3, that this joint restriction seems to hold prior to 1981, while - with a few exceptions - it is violated throughout the rest of the sample period.

[17] In an earlier draft of this paper, I also employed the multivariate Beveridge-Nelson decomposition from a VEqCM where the balanced budget restrictions was imposed, in an effort to identify whether the deviations of expenditures and revenues from a balanced budget was due to ‘trend’ or due to ‘cyclical’ movements. The ‘cyclical’ components from that exercise were not stationary. I obtained similar results employing the Gonzalo and Granger (1995) decomposition and its modification by Proietti (1997). Therefore the distinction between ‘trend’ and ‘cyclical’ movements in spending and revenue was inadequate in identifying whether spending or revenues were more responsible for the fiscal deficit. I am grateful to an anonymous referee of this journal pointing this out to me.

[18] I am grateful to an anonymous referee for bringing this to my attention.



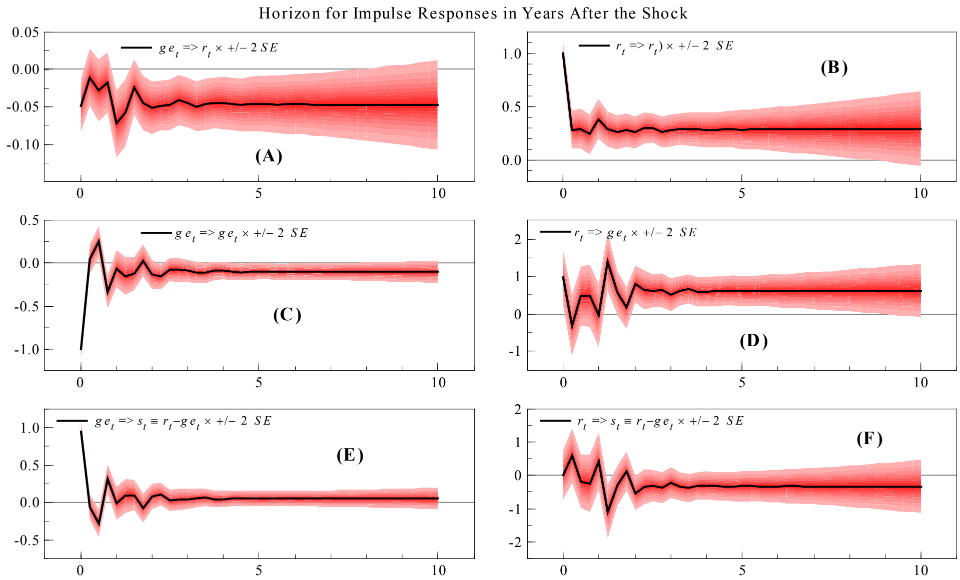
Notes: Panel A reports the recursively-estimated LR test for intertemporal solvency, Panel B the recursively calculated LR test of equal adjustment coefficients, Panel C the recursive estimates of β_{ge} under the assumption that $\alpha_r = \alpha_{ge}$, and Panel D the *LR* test of intertemporal solvency and equal adjustment coefficients.

Figure 3. Recursively Calculated Neutrality Tests

Unfortunately, the neutrality test is only a limited tool for assessing the efficiency of alternative strategies of balancing a budget through cuts in expenditures or through tax increases, because it is informative only for the long-run, whereas the adjustment path during the transition period may matter considerably in relation to political issues, owing to the relatively short-term political perspective in representative democracies. For this reason, I have also employed impulse response functions associated with the VEqCM summarized in Table 3. Instead trying to identify "structural" shocks, here I proceed using GIRs.[19] Such an exercise provides valuable insights about the long-run change of

[19] In an earlier draft of the paper I have assumed that a discretionary cut in expenditure is defined as that part of innovation in expenditures, which is orthogonal to revenue innovation and similarly, a discretionary revenue increase aiming at deficit reduction has to be orthogonal to spending innovations. Here I would like to thank an anonymous referee of this journal for pointing out that, it is hard to argue that discretionary revenue (expenditure) innovations can be derived form a VEqCM with the whole revenue (expenditure) innovations.

both the expenditure and revenue ratio, i.e., of the ‘size of government’. The graphical output of the results is summarized in Figure 4.



Notes: The confidence intervals calculated by Monte Carlo Simulation. Panels (A), (C), (E) display the effects of a (generalized) government expenditure shock normalized to correspond to a - 1% decrease to r_t (Panel A), to ge_t (panel C) and to $s_t = r_t - ge_t$. Panels (B), (D) and (F) display the effects of an increase in revenue normalized to correspond to 1%.

Figure 4. Generalized Impulse Responses

Panels A, C and E of Figure 4 report the impulse responses to unit expenditure cut shock, whereas panels B, D, and F the impulse responses to a unit increase in government revenues. First notice that the generalized nature of these impulse responses is apparent from the simultaneous behavior of the both variables in the system. Typically, a positive shock to revenue is associated with a positive shock to spending, and *vice versa*. Additionally, what is remarkable in these plots is that the previous analysis is confirmed and strengthened. Observe that that a 1% cut in expenditure leads to permanently lower expenditure and revenue levels, but it also leads to an improvement of the budget surplus, with the long-run effect being 0.053%. On the other hand a 1% increase in revenue, leads to permanently higher level of revenue, but government expenditure is increased by much more, leading to a worsening of the budget surplus of the order of magnitude of 0.333%. Remarkably, the impulse responses also indicate that the Greek fiscal authorities behave according to the *tax and spend hypothesis*, since as discussed above, in this case an increase in government revenues would result in even higher government deficit. Another important feature that comes forth is that reduction of the government deficit can be more efficiently performed via reductions in spending rather than increasing taxation.

6. CONCLUDING REMARKS

This study has used some recent advances in the testing of long-run equilibrium relationships to shed light on the long-run relation between the two Greek government's budget components. More specifically, the aim was to describe the historical behavior of the Greek government and to suggest ways of eliminating the budget deficit by analyzing the relative efficiency of *revenue increase* vs. *spending reduction*. Several recent developments in the econometric analysis of non-stationary processes and cointegration were applied and a number of novel results stem from the analysis.

First, using Johansen's (1995) FIML testing strategy, it was established that government spending and revenues share a common stochastic trend, but the restriction for intertemporal solvency is violated for the whole sample examined. I then examined the parameter stability of the cointegrating relation estimated by means of tests proposed by Hansen and Johansen (1999) and my results indicated that the cointegrating space is stable over the recursive sub-sample period and that the estimated coefficients do not exhibit instabilities in recursive parameter estimates.

Second, utilizing causality tests, I found that revenue precedes government expenditures and that the behavior of the Greek fiscal authorities is consistent with the *tax and spend hypothesis*. I have further evaluated the relative efficiency of fiscal adjustments via *revenue* or *spending*, showing that the two alternative fiscal stabilization measures are not neutral. More specifically, recursive tests showed that the neutrality hypothesis might have been valid for some periods in the sample, especially before the 1990s, but post 1990 it is clearly violated.

Finally, in an effort to examine the dynamic interaction of the Greek fiscal variables, I have utilized generalized impulse response analysis. This also indicated that raising revenue would lead to higher spending, thus worsening the budget balance, whereas decreasing spending would lead to an improvement of the deficit. Therefore, the policy implications of the analysis seem quite clear. In order for the government to restore its intertemporal budget balance, government expenditures must begin to grow at a lower rate than that at which real revenues must rise.

REFERENCES

Akaike, H. (1969), "Fitting Autoregressive Models for Prediction," *Annals of the Institute of Statistical Mathematics,* 21, 243-7.

Barro, R.L. (1979), "On the Determination of Public Debt," *Journal of Political Economy,* 87, 940-961.

Beveridge, S., and C.R. Nelson (1981), "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the Business Cycle," *Journal of Monetary Economics*, 7, 151-74.

Bohn, H. (1991), "Budget Balance Through Revenue or Spending Adjustments?" *Journal of Monetary Economics*, 27, 333-359.

Dickey, D., and W. Fuller (1981), "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root," *Econometrica*, 49, 1057-1072.

Dolado, J. (1992), "A Note on Weak Exogeneity in VAR Cointegrated Models," *Economics Letters*, 38, 139-43.

Doornik, J. (1998), "Approximations to the asymptotic distribution of cointegration tests," *Journal of Economic Surveys*, 12, 573-593.

Elliott, G., T. Rothenberg, and J. Stock (1996), "Efficient Tests of an Autoregressive Unit Root," *Econometrica*, 64, 813-836.

Friedman, M. (1978), "The Limitations of Tax Limitation," *Policy Review (Summer),* 7-14.

Garcia, S., and P.-Y. Henin (1999), "Balancing Budget through Tax Increases or Expenditure Cuts: Is it Netral?" *Economic Modelling,* 16, 591-612.

Gonzalo, J. (1994), "Five Alternative Methods of Estimating Long-Run Equilibrium Relationships," *Journal of Econometrics*, 60, 203-223.

Gonzalo, J., and C.W.J. Granger (1995), "Estimation of Common Long-Memory Components in Cointegrated Systems," *Journal of Business and Economic Statistics*, 13, 27-35.

Granger, C.W.J. (1969), "Investigating Casual Relations by Econometric Models and Cross-Spectral Methods," *Econometrica,* 37, 424-438.

_____ (1988), "Causality, Cointegration and Control," *Journal of Economic Dynamics and Control*, 12.

Hakkio, C.S., and M. Rush (1991), "Is the Budget Deficit 'Too Large'?" *Economic Inquiry*, 29, 429-45.

Hamilton, J.D., and M.A. Flavin (1986), "On the Limitations of Government Borrowing: A Framework for Testing," *American Economic Review*, 76, 808-819.

Hansen, H. and K. Juselius (1995), *CATS for RATS*, Evanston, IL: Estima.

_____ and S. Johansen (1999), "Some Tests for Parameter Constancy in Cointegrated VAR-Models," *Econometrics Journal,* 2, 306-333.

Haug, A. (1995), "Has Federal Budget Deficit Policy Changed in Recent Years?" *Economic Inquiry*, 33, 104-118.

Hondroyiannis, G., and E. Papapetrou (1996), "An Examination of the causal relationship between government spending and revenue: a cointegration analysis," *Public Choice*, 89, 363-374.

Hoover, K.D., and S. Sheffrin (1992), "Causation, Spending and Taxes: Sand in the Sandbox or Tax Collector for the Welfare State?" *American Economic Review*, 82, 225-248.

Johansen, S. (1995), *Likelihood-Based Inference in Cointegrated Vector-Autoregressive Models,* Oxford: Oxford University Press.

Kollias, C., and S. Makrydakis (2000), "Tax and Spend or Spend and Tax? Empirical Evidence from Greece, Spain, Portugal and Ireland," *Applied Economics,* 32, 533-546.

Koop, G., H. Pesaran, and S. Potter (1996), "Impulse Response Analysis in Nonlinear Multivariate Models," *Journal of Econometrics*, 74, 119-147.

Koren, S., and A. Stiassny (1998), "Tax and Spend, or Spend and Tax? An International Study," *Journal of Policy Modeling*, 20, 163-191.

Kremers, J.M. (1988), "Long-Run Limits on U.S. Federal Debt," *Economics Letters*, 28, 259-62.

_____ (1989), "U.S. Federal Indebtedness and the Conduct of Fiscal Policy," *Journal of Monetary Economics*, 23, 219-238.

Kwiatkowsky, D., P.C.B. Phillips P. Schmidt, and Y. Shin (1992), "Testing the Null of Stationarity Against the Alternative of a Unit Root: How Sure Are We that Economic Series Have a Unit Root?" *Journal of Econometrics,* 54, 159-178.

Lütkepohl, H., and H. Reimers (1992), "Impulse Response Analysis of Cointegrated Systems," *Journal of Economic Dynamics and Control*, 16, 53-78.

Newey, W., and K. West (1987), "A Simple, Positive Semi-definite, Heteroscedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica,* 55, 703-708.

Ng, S., and P. Perron (2001), "Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power," *Econometrica*, 69, 1519-1554.

Owoye, O. (1995), "The Causal Relationship Between Taxes and Expenditures in the G7: Cointegration and Error-Correction Models," *Applied Economics Letters,* 2, 19-22.

Peacock, A.T., and J. Wiseman (1979), "Approaches to the Analysis of Government Expenditure Growth," *Public Finance Quarterly,* 7, 3-23.

Pesaran, H.M., and Y. Shin (1998), "Generalized Impulse Response Analysis in Linear Multivariate Models," *Economics Letters*, 58, 17-29.

Proietti, T. (1997), "Short-Run Dynamics in Cointegrated Systems," *Oxford Bullettin of Economics and Statistics*, 59, 405-422.

Quintos, C.E. (1995), "Sustainability of the Deficit Process with Structural Shifts," *Journal of Business and Economic Statistics*, 13, 409-417.

Ram, R. (1988), "Additional Evidence on Causality Between Government Revenue and Government Expenditure," *Southern Economic Journal,* 54, 763-769.

Ram, R., J.H. Stock, and M.W. Watson (1990), "Inference in Linear Time Series Models with Some Unit Roots," *Econometrica*, 58, 113-144.

Sims, C.A. (1980), "Macroeconomics and Reality," *Econometrica*, 48, 1-48.

Stock, J.H. and M.W. Watson (1988), "Testing for Common Trends," *Journal of the American Statistical Association*, 83, 1097-1107.

Trehan, B., and C.E. Walsh (1988), "Common Trends, the Government Budget Constraint, and Revenue Smoothing," *Journal of Economic Dynamics and Control*, 12, 425-44.

_____ (1991), "Testing Intertemporal Budget Constraints: Theory and Applications to U.S. Federal Budget and Current Account Deficits," *Journal of Money, Credit, and Banking*, May 1991, 206-23.

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