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The Causal Relationship between Tax Revenues and Government Spending in Greece: 1950-1990

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ABSTRACT

The paper investigates the direction of causality between tax revenues and government spending in the case of Greece. The empirical methodology employed is that of cointegration and error correction which provides additional channels through which causality between two variables can be established. The findings reveal unidirectional causality from government spending to tax revenues which suggests that tax and spending decisions are not made jointly by the Greek fiscal authorities.

1. INTRODUCTION

The steady expansion of the public sector in many developed economies has attracted the interest of many theoretical and empirical economists (Peltzman, 1980; Meltzer and Richard, 1981). A number of hypotheses have been advanced and empirically tested with varying degrees of success in an attempt to explain this phenomenon. In this context, the relationship between government expenditures and revenues has received considerable attention in the literature. The usual approach to evaluating this relationship rests with the deployment of standard Granger causality tests (Anderson et al., 1986; Manage and Marlow, 1986; Ram, 1988; Provopoulos and Zambaras, 1991). Recently, however, this practice has been questioned as neglecting the presence of an additional channel of causation that exists when tax revenues and public expenditure happen to share a common trend. In this case, it is argued, temporal causality should be investigated through the application of appropriately defined error-correction specifications otherwise misleading inferences may result (Miller and Russek, 1990; Owoye, 1995).

Overall, the empirical findings reported in the literature are not consistent. Bidirectional as well as unidirectional causality between government expenditures and revenues has been documented (Anderson et al., 1986; Manage and Marlow, 1986; Ram, 1988; Owoye, 1995; Belessiotis, 1995). The bulk of the empirical work undertaken so far, concentrates almost exclusively on the USA and the developed countries of the G7 group. The scope of this study is to investigate empirially the causal relationship between government spending and tax revenues in the case of Greece. The causality link between revenues and spending in Greece has also been examined by Provopoulos and Zambaras (1991) using the standard Granger test for the period 1957-1987 and the reported results point to unidirectional causality from spending to revenues. In the present study the sample period covers the years 1950-1990 which have seen important socioeconomic developments (Jouganatos, 1992; Alogoskoufis, 1995). Throughout this period, successive Greek governments have played a key role in the country's post-war development with the degree of state intervention in the economy steadily expanding over the years. This is particularly true for the late 70s and the 80s when extensive nationalisation programmes took place and unsustainable deficit finance practices prevailed leading to sharp rises in government debt (Courakis et al., 1993; Alogoskoufis, 1995, Makrydakis et al., 1995). Belessiotis (1995) notes that one of the key convergence criteria for the participation of Greece to the European Monetary Union concerns the correction of fiscal imbalances. In this context it is of importance to establish the principle cause of fiscal deficits by examining the interdependence between spending and revenues. In order to determine the causal relationship between government spending and revenues in Greece, we follow the recent trend of employing integration and cointegration testing techniques and the related notion of error-correction modelling (Perman, 1991; Holden and Thompson, 1992). In this way the pitfalls associated with the application of standard Granger causality tests can be avoided.

2. GOVERNMENT SPENDING AND TAX-REVENUES IN GREECE

The relationship between government expenditures and revenues has, as already been noted, given rise to number of empirical studies. Researchers have attempted to establish empirically whether government expenditures cause government tax revenues or vice versa. Establishing the direction of interdependence between these two macroeconomic aggregates could assist policymakers in identifying the source of any fiscal imbalances that might exist and in designing a suitable fiscal reform. Thus, if on the one hand, bidirectional causality is established this would be in accordance with the fiscal synchronisation hypothesis which implies that tax and spending decisions are jointly made by the fiscal authorities. If, on the other hand, unidirectional causality is detected this is taken to be an indication that spending decisions are based on tax revenues or the opposite. In economic theory these cases are known as the tax-and-spend hypothesis where tax revenues lead to changes in government spending or the spend-and-tax hypothesis where spending leads to changes in tax revenues (Anderson et al., 1986).

Manage and Marlow (1986) report unidirectional causality from taxes to expenditures in the case of the United States for the period 1929-1982. On the other hand, Anderson et al. (1986) have concluded that the reverse is the case, that is government expenditures Granger-cause government taxes. Owoye (1995) reports bidirectional causality in five of the G7 countries and unidirectional, from taxes to expenditures, in the case of Japan and Italy.

In the case of Greece for the years 1950-1990 both government expenditures and revenues have displayed sharp upward trends (see, Figure 1). Government spending stood at 19.9% of GDP in 1950 and by 1990 it has reached 50.5% (see, Table 1). Courakis et al (1993) note that public spending is the outcome of a complex process reflecting economic priorities, institutional factors and the ideological proclivities of the governments. In the case of the Greek economy, a large part of this rise can also be attributed to high levels of defence spending (Kollias, 1995) as well as to extensive nationalisation programmes and substantial increases in public sector employment (Courakis et al., 1993).

More specifically, Karavitis (1987) argues that developmental aspects of the economy, institutional features and demographic factors have played a significant part in the growth of government expenditures in Greece during the post-war period. Georgakopoulos and Loizides (1994) using Wagner type models test alternative hypotheses of the growth of government size in Greece. They conclude that industrialisation, urbanisation and population changes have played a considerable role in the growth of government expenditures.

FIGURE 1

Government expenditures and tax revenue as a share of GDP

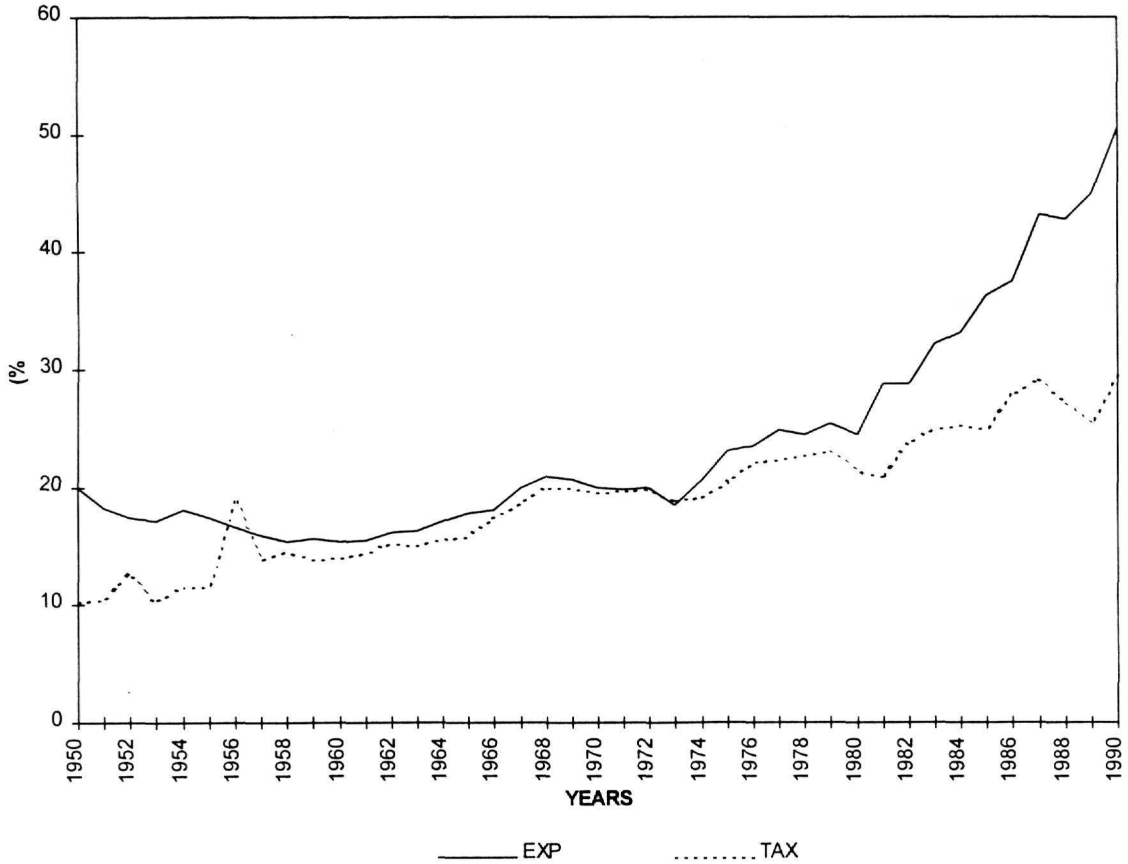

TABLE 1

Government spending and tax revenues as a share of GDP

	1950	1955	1960	1965	1970	1975	1980	1985	1990
exp/gdp	19.9	17.4	15.4	17.7	20	23.1	24.4	36.2	50.5
tax/gdp	10.4	11.6	13	15.7	19.6	20.4	21.3	24.8	29.3

Source: Bank of Greece (1992).

Similar upward trends have been exhibited by government tax revenues which have increased from about 10.4% of GDP in 1950 to 29.3% in 1990 (see. Table 1).

Given the preceding discussion, it would be interesting to examine what has caused what, that is to determine the direction of causality between tax-revenues and expenditures in the case of Greece. Evidence of one-way causality from revenues to expenditures would imply that higher revenues have led to higher spending, while causality in the opposite direction would suggest that governments in the case of Greece decide how much to spend and then look for revenue sources to finance that level of spending. This can be achieved through both direct and indirect taxation, borrowing and money creation. Finally, bidirectional causality would indicate that simultaneity between revenues and spending decisions is exhibited by the Greek fiscal authorities.

3. GRANGER CAUSALITY AND COINTEGRATION

The notion of causal relationships is the basis on which the empirical testing of theoretical propositions rests. The most commonly used technique in establishing the direction of causality between two variables, say X_t and Y_t, is the Granger causality test. In formal terms this test is based on the estimation of the following regression:

$$\Delta X_t = \pi_0 + \sum_{i=1}^{p} \pi_{xi} \Delta X_{t-i} + \sum_{i=1}^{p} \pi_{yi} \Delta Y_{t-i} + \psi_t \tag{1}$$

where Δ is the first difference operator, p is the maximum lag length, ΔX_t, ΔY_t are stationary time series and X_t and Y_t are not cointegrated. If Ω_t is a universe of information up to and including period t then the Granger (1969) definition of causality states that Y causes X, given Ω_t, if current changes of X can be predicted better using past changes of Y (i.e. ΔY_s, $s < t$) than by not using them at all. Thus, if past ΔY contributes significantly to forecasting current ΔX, that is the coefficients π_{yi}, $i=1, 2,...,p$, in (1) are jointly significant by the standard F-test, then Y Granger causes X. Similarly, if ΔY_t replaces ΔX_t as the dependent variable in (1) then the null hypothesis that X does not Granger cause Y is rejected if π_{xi}, $i=1, 2,...,p$, are jointly significant. The possible outcomes that one can reach by using Granger causality tests for two variables X and Y are: a) neither variable Granger causes the other, b) Y causes X but not the opposite, c) X causes Y but not vice versa and d) both variables Granger cause each other.

Although the standard Granger causality test may be generally sufficient for determining the direction of causality, it can lead to the drawing of misleading inferences when the variables under study are integrated of the same order and share a common trend, that is they are cointegrated according to the definition of Engle and Granger (1987). As Miller and Russek (1990), Owoye (1995), and Demetriades and Hussein (1996), point out, the presence of cointegration between X and Y identifies an additional channel of causation which will go undetected had the standard Granger causality test is employed. In this case the utilisation of equation (1) simply fails to account for any long-run information embodied in the data. They suggest therefore that, in the event of cointegrating variables, the direction of causality should be investigated by means of the following set of regressions:

$$\Delta X_t = \pi_0 + \sum_{i=1}^{p} \pi_{xi} \Delta X_{t-i} + \sum_{i=1}^{p} \pi_{yi} \Delta Y_{t-i} + \phi \varepsilon_{t-1} + \psi_{1t} \tag{2}$$

$$\Delta Y_t = \pi_0 + \sum_{i=1}^{p} \pi_{xi} \Delta X_{t-i} + \sum_{i=1}^{p} \pi_{yi} \Delta Y_{t-i} + \vartheta \varepsilon_{t-1} + \psi_{2t} \tag{3}$$

where ε_{t-1} is the lagged error term from the standard cointegrating equation:

$$X_t = \alpha + \beta Y_t + \varepsilon_t \quad (4)$$

Equations (2) and (3) have the familiar error-correction model (ECM thereafter) specification and state that Y (X) could Granger cause X (Y) even if past changes of Y (X) do not contribute to forecasting current changes of X (Y). All that is needed for determining the direction of causality in this case is the existence of a common trend between X and Y. If such a common trend exists, that is X and Y are cointegrated, then the error term ε_{t-1} in (2) or (3) should be statistically significant establishing the additional line of causation mentioned above.

4. THE EMPIRICAL FINDINGS

The direction of causality between government spending and tax revenue in Greece is investigated using annual data covering the period 1950-1990. The data for the variables in question are obtained from the *Long-Run Statistical Series of the Greek Economy*, Bank of Greece, 1992, Tables 31 and 32 respectively. For estimation purposes the variables are logarithmicly transformed and will be denoted henceforth by $lexp_t$ and $ltax_t$ respectively.

Following the standard practice each variable is first tested for the presence of unit roots in its autoregressive representation. If both series are found to be of the same order of integration a cointegrating regression like (4) is run and cointegration is evaluated. If the presence of a long-run relationship between $lexp_t$ and $ltax_t$ is established then two error correction specifications based on equations (2) and (3) are estimated. Finally, three types of tests are run in order to determine the direction of causality between the variables of interest. First, a conventional Granger causality F-type test on the joint significance of the lagged differenced terms in equations (2) and (3) is computed (that is, ΔY_{t-i} in (2) and ΔX_{t-i} in (3)). Second, the significance of the lagged error correction term in either ECM specification is evaluated through the use of a t-ratio test. Third, F-tests are calculated to assess the joint significance of the lagged difference terms and the lagged cointegrating vector in each ECM specification.

4.1. Testing for the order of integration

The historical series on government spending and tax revenue are examined for their order of integration by means of the standard Augmented Dickey-Fuller (ADF) testing principle (Said and Dickey, 1984). The computation of the ADF test statistics for a variable Z_t, is based on the application of the following auxiliary regression:

$$Z_t = \gamma + \delta t + \rho Z_{t-1} + \sum_{i=1}^{k} c_i \Delta Z_{t-i} + \xi_t \tag{5}$$

where ξ_t is a sequence of normal, independent random variables with mean 0 and variance σ^2. Expression (5) permits the evaluation of the most general alternative hypothesis that of stationarity around a deterministic trend. Moreover, the inclusion of the deterministic trend in (5) serves the additional purpose of rendering the unit root test invariant to the value of the drift term (Kiviet and Phillips, 1992). The maximum lag length k adopted for correcting serial correlation in the associated ADF auxiliary regressions is determined on the basis of evidence provided by sequential t-ratio tests on the significance of the highest order lag in the estimated autoregression coupled with extensive residual diagnostic testing (Ng and Perron, 1995). In addition to the ADF tests, we have also evaluated directly the hypothesis of stationarity by implementing the recently developed η_{μ} and η_{τ} tests of Kwiatkowski, Phillips, Schmidt, and Shin (1992), or in terms of their initials KPSS thereafter.[1] The KPSS semi-parametric procedure tests for level (η_{μ}) or trend stationarity (η_{τ}) against the alternative of a unit root and its combined use with the ADF test is likely to enable more clear-cut conclusions to be drawn with regard to the order of integration of the series under investigation given the small size of our sample. Finally, the first differences of the relevant variables are subjected to the same battery of tests in accordance to the suggestions of Dickey and Pantula (1987) so that the presence of higher order integrated processes can also be examined.

Since the tests are conducted under the null hypothesis of a unit root and under the null hypothesis of stationarity, four are the possible outcomes that can emerge from the joint implementation of integratibility tests, i.e.:

a) If the null of trend stationarity is accepted and that of a unit root is rejected then it can be concluded that the series under investigation is trend stationary.
b) If the null of trend stationarity is rejected and the null of a unit root is accepted then the series is non-stationary.
c) If both null hypotheses are accepted then there is insufficient information to deter-mine whether there is stationarity or not.
d) If both nulls are rejected then no definitive conclusion on the order of integration of the series involved can be reached.

Table 2 summarises the results of the ADF and KPSS testing procedures for $lexp_t$ and $ltax_t$. As far as the levels of the variables are concerned, the ADF test assumes values which do not exceed the 5% critical values suggesting the presence of at least one unit root. This finding is also corroborated by the KPSS tests that are found to reject the null hypothesis of (trend) stationarity at the same level of significance. When the same testing procedures are applied to the first differences of the series, the presence of a second unit root is not supported by either procedure at the 95% confidence level. Hence the statistical evidence points to the

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[1]. The KPSS test statistic for level or trend stationarity is given by:

$$\eta = \frac{1}{s^2(k)T^2}\sum_{t=1}^{T} S_t^2$$

where $S_t = \sum_{i=1}^{t} u_i$, u_t are the residuals from the regression of X_t on a constant or a constant and a trend for level or trend stationarity respectively, $s^2(k)$ is the non-parametric estimate of the "long-run variance" of u_t, and k stands for the lag truncation parameter.

conclusion of both variables being described by I(1) processes. Having established that $lexp_t$ and $ltax_t$ are integrated of the same order, we can now proceed with testing whether the series in question move together in the long-run, that is whether they share a common trend.

TABLE 2
Unit root and stationary tests

Variable	Dickey-Fuller tests		KPSS tests	
	t_{adf}	k	η_μ	η_τ
$lexp_t$	-0.387	1	1.096**	0.284**
$\Delta lexp_t$	-3.974*	2	0.846**	0.066
$ltax_t$	-1.169	0	1.091**	0.257**
$\Delta ltax_t$	-7.915**	0	0.321	0.078

Notes:
1. * and ** indicate significance at 5% and 1% respectively.
2. η_μ , and η_τ stand for the KPSS tests for level and trend stationarity respectively. The l(4) formula of Schwert (1987) is used for the determination of the lag truncation parameter. The critical values for these tests are obtained from the response surface estimates of KPSS critical values that appear in Table 1 in Sephton (1995).
3. k represents the number of lagged differences required to account for serial correlation in the ADF auxiliary regressions.
4. The ADF auxiliary regressions are fitted with both a constant and a trend. The critical values for the t_{adf} statistics are taken from the relevant response surface estimates of Table 1 in MacKinnon (1991).

4.2. Testing for cointegration

Following Engle and Granger (1987) the existence of cointegration between $lexp_t$ and $ltax_t$ is examined by subjecting to unit root testing the residuals from the following regressions:

$$ltax_t = \alpha_0 + \beta lexp_t + \varepsilon_{1t} \quad (6)$$

$$ltax_t = \alpha_0 + \alpha_1 t + \beta lexp_t + \varepsilon_{2t} \quad (7)$$

The estimation of both of the above expressions intends to help us distinguish between the notions of deterministic and stochastic cointegration for the series under consideration. Specifically, if cointegration is established by means of the *demeaned* specification (6) then this may correspond to deterministic cointegration, which implies that the estimated cointegrating vector eliminates both stochastic and deterministic trends. If, on the other hand, cointegration is detected through the estimation of the *detrended* expression (7), that is the linear stationary combinations of the I(1) variables at issue have non-zero linear trends, this will correspond to the notion of stochastic cointegration.[2]

In contrast to the usual practice, equations (6) and (7) have not been estimated by OLS because this technique yields estimates of the cointegrating vector which, in spite of their superconsistency, are inefficient, while statistical inference on the cointegrating vector is, in general, non-standard (see, Ogaki, 1993). Alternative asymptotically efficient estimation techniques for cointegrating relationships that alleviate most of the deficiencies of OLS have been proposed in the literature. The most commonly used one in empirical applications is that of Johansen (1991). However, its implementation in small samples, as the one used in the present study, is inadvisable since the Johansen estimator is subject to significant small sample bias (Phillips, 1994). Furthermore, the utilisation of a single equation estimation method instead of the Johansen system estimation technique is more appropriate in our case because with two variables there can only be one cointegrating vector. Estimators suitable for single equation estimation which exhibit better finite sample performance to that of the Johansen estimator are the Phillips-Hansen Fully Modified (FM) OLS estimator (Phillips and Hansen, 1990) and the DOLS estimator of Saikkonen (1991) and Stock and Watson (1993). In the present study the demeaned and detrended cointegrating regressions have been estimated through the implementation of the Phillips-Hansen FM estimator. Our preference of FM over DOLS is based on the fact that the FM estimator corrects for serial correlation and endogeneity semi-parametrically without reducing the size of the sample available for estimation in contrast to the use of the parametric leads and lags correction employed by DOLS.

In order to determine whether $lexp_t$ and $ltax_t$ are stochastically or deterministically cointegrated, the residuals obtained from the FM estimated demeaned and detrended regressions are subsequently subjected to integration testing. In particular, the null hypothesis of no cointegration between the two series is investigated by applying the ADF testing principle. Naturally, the estimated ADF auxiliary regressions do not involve any deterministic components as the latter have been accounted for in the relevant cointegrating regressions (Engle and Yoo, 1987). The resultant ADF t statistics for the residuals of equations (6) and (7) are denoted by t^{μ}_{adf} and t^{τ}_{adf} respectively. Moreover, the decision on the number of lagged differenced terms in the ADF auxiliary regressions has been made on the basis of the same criterion as in the case of integration testing. Following the methodological principles of the previous section, the null of cointegration, i.e. the null hypothesis of stationarity, is also examined by computing appropriately defined KPSS test statistics (see, Shin, 1995). The KPSS test statistic calculated from the residuals of the demeaned regression is denoted by C_{μ} and is appropriate for testing the null of deterministic

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2. Ogaki, 1993, provides a lucid presentation of the definitions of stochastic and deterministic cointegration.

cointegration, while the null hypothesis of stochastic cointegration is evaluated by means of the KPSS statistic C_τ obtained from the residuals of the detrended regression.[3]

The Phillips-Hansen FM OLS estimated cointegrating regressions along with the conducted cointegration and no cointegration tests are reported in Table 3. Given that the principle of deterministic cointegration is stronger than the concept of stochastic cointegration we examine first the evidence from the detrended regression (7). To put it differently, we conduct testing from the most general to the most restrictive hypothesis. The findings of Table 3 on the detrended regression provide strong evidence of stochastic cointegration between government spending and tax revenues. This is so because the null hypothesis of cointegration cannot be rejected, even when the 10% critical values for $C\tau$ are considered, while the null of no cointegration is soundly rejected at the 1% level of significance. If we now turn to the results of the demeaned regression, no clear conclusion can be arrived at with regard to the presence of deterministic cointegration between $lexp_t$ and $ltax_t$. The t^{μ}_{adf} statistic assumes the value of -2.145 suggesting that the null hypothesis of no cointegration cannot be rejected, while the C_μ statistic is found to be equal to 0.279 indicating that the null of deterministic cointegration cannot be rejected either at the usual 95% confidence. Since the existence of stochastic cointegration has been firmly founded between the variables of interest, it is now possible to proceed with the setting up of the relevant error correction specifications on the basis of which the direction of causality can be determined.

[3]. For a detailed discussion on the formulation of KPSS cointegration tests the reader is referred to Shin, 1995.

TABLE 3

Phillips-Hansen fully modified estimates and cointegration tests

Cointegrating regression: $ltax_t = \alpha_0 + \alpha_1 t + \beta lexp_t + \varepsilon_t$					
	α_0	α_1	β	C_μ or C_τ	t^μ_{adf} or t^τ_{adf}
Demeaned	-0.041 (0.249)		0.988 (0.021)	0.279	-2.145 [1]
Detrended	2.801 (0.328)	0.059 (0.007)	0.633 (0.041)	0.056	-5.152** [0]

Notes:

1. * and ** indicate significance at 5% and 1% respectively.
2. C_μ , and C_τ stand for the KPSS cointegration tests when the cointegrating regression involves a constant and a constant and a trend respectively. The l(4) formula of Schwert (1987) is used for the determination of the lag truncation parameter. The critical values for these tests are obtained from Table 1 in Shin (1994).
3. The t^μ_{adf} and t^τ_{adf} statistics are computed on the basis of cointegrating regressions involving a constant and a constant and a trend respectively. The critical values for the t_{adf} statistics are taken from the relevant response surface estimates of Table 1 in MacKinnon (1991).
4. The figures in parentheses are standard errors while the numbers in brackets correspond to the number of lags in the ADF auxiliary regressions.

4.3. Testing for Granger causality

In accordance to the discussion of Section 3, the ECM specifications (2) and (3) were estimated by OLS with $ltax_t$ and $lexp_t$ standing for X_t and Y_t respectively and the error correcting term being formulated from the residuals of equation (7). The maximum lag length p has been determined by sequential F-type tests on the joint significance of the highest order lagged differenced terms and comprehensive diagnostic testing. Given the annual frequency of the data set, a maximum lag length of 2 was initially chosen for both equations. Eventually, though the testing procedure revealed that the ECM regression for $ltax_t$ required a lag length of one while two lags were needed for the $lexp_t$ error correction specification. The estimated ECM equations were subsequently used as the platform for assessing the null hypothesis of no causation according to the guidelines given at the beginning of Section 4. This hypothesis is evaluated by a t-ratio and two F-type tests. The results of the OLS estimation along with the computed causality tests for both equations are presented in Table 4.

TABLE 4

Error correction models based on the cointegrating regression of tax revenue on public expenditure and tests for temporal causality

Regressant	Regressors						Causality tests	
	$\Delta ltax_{t-1}$	$\Delta ltax_{t-2}$	$\Delta lexp_{t-1}$	$\Delta lexp_{t-2}$	ecm_{t-1}	Constant	Standard	ecm test
$\Delta ltax_t$	-0.004 (0.155)		0.559 (0.19)		-0.87 (0.209)	0.076 (0.034)	F(1,34) 8.624**	t ratio -4.167**
	Diagnostics: R^2=0.55, S.E.=0.088, AR: F(2,32)=1.839, ARCH: F(1,32)=0.024, NORM: $\chi^2(2)$=5.98, RESET: F(1,33)=0.178						Joint test F(2,34)=18.76**	
	$\Delta ltax_{t-1}$	$\Delta ltax_{t-2}$	$\Delta lexp_{t-1}$	$\Delta lexp_{t-2}$	ecm_{t-1}	Constant	Standard	ecm test
$\Delta lexp_t$	-0.102 (0.143)	-0.148 (0.102)	0.554 (0.17)	0.416 (0.15)	-0.049 (0.169)	0.049 (0.025)	F(2,31) 1.086	t ratio -0.289
	Diagnostics: R^2=0.59, S.E.=0.055, AR: F(2,29)=1.203, ARCH: F(1,29)=0.10, NORM: $\chi^2(2)$=5.31, RESET: F(1,30)=3.587						Joint test F(3,31)=1.402	

Notes: :

1. * and ** indicate significance at 5% and 1% respectively.
2. The standard Granger causality test is an F test evaluating the joint significance of the differenced lagged revenue (or expenditure) terms in the ecm specification when the regressant is the diferrenced expenditure (revenue) variable.
3. The ecm test is the usual t-test assessing the significance of the ecm term alone in the models considered.
4. The joint test is an F test assessing the joint significance of the differenced lagged revenue (expenditure) terms and the lagged ecm term when the dependent variable in the ecm specification is the diferrenced expenditure (revenue) variable.
5. S.E. is the equation standard error, AR is the LM test for serial correlation, ARCH tests for conditional heteroscedasticity, NORM is the Jarque-Berra test for normality, and RESET tests for the validity of the specification.
6. The figures in parentheses are standard errors.

The empirical findings suggest unequivocally the existence of unidirectional causality from government spending to tax revenue. Not only the lagged error correction term and the lagged $\Delta lexp_t$ in equation (2) are individually significant at the 99% confidence level but also the null hypothesis H_0: $\pi_{lexp1}=0$ and $\varphi=0$ is soundly rejected as the relevant F-test clearly indicates (see, Table 4). On the contrary, causality from tax revenue to government spending is not substantiated by the evidence. All three types of causality tests employed on equation (3) are found to be insignificant at both the 5% and 1% levels of significance. These results point to the conclusion that the Greek fiscal authorities make their tax and spending decisions independently of one another with the latter taking precedence over the former.

5. CONCLUSIONS

The scope of this paper was to investigate empirically the causal relationship between tax-revenues and government spending in the case of Greece for the period 1950-1990. To this end the notion of cointegration and error correction modelling have been used as a basis for inference. The results obtained identify the existence of unidirectional causality from government expenditures to tax-revenues while no evidence of any feedback was found. The same direction of causality has also been reported by Provopoulos and Zambaras (1991) albeit for a shorter time period using the standard Granger test. Given the fact that correction of fiscal imbalances constitutes an important convergence criterion for participation of Greece in the EMU, a policy implication of the findings reported herein is that the necessary fiscal adjustment should principally be achieved through the curtailing of government spending.

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