# The nexus between public expenditure and inflation in the Mediterranean countries

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2011

# ***THE NEXUS BETWEEN PUBLIC EXPENDITURE AND INFLATION IN THE MEDITERRANEAN COUNTRIES***

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*Abstract: The aim of this article is to assess the empirical evidence of the nexus between public expenditure and inflation for the Mediterranean countries during the period 1970-2009, using a time-series approach. After a brief introduction, a concise survey of the economic literature on this issue is shown, before discussing the data and introducing some econometric techniques. Stationarity tests reveal, generally, that public expenditure/GDP ratio is a I(1) process, while prices index is a I(2) process. Moreover, we find a long-run relationship between the growth of public expenditure and inflation only for Portugal. Furthermore, Granger causality tests results show a short-run evidence of a directional flow from expenditure to inflation for Cyprus, Malta and Spain; of a bidirectional flow for Italy; and from inflation to public expenditure for France. Some notes on the policy implications of our empirical results conclude the paper.*

*Summary: 1. Introduction; 2. Literature Survey; 3. Data and methodology; 4. Econometric results; 5. Conclusions and policy implications.*

*Keywords: public expenditure; inflation; time-series; unit root; cointegration; causality; Mediterranean countries.*

*JEL Codes: C32; E31; E62; H50; N44.*

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*Acknowledgements:* Comments from Francesco Forte, Roberto Golinelli and Gian Cesare Romagnoli are gratefully acknowledged. All remaining errors are entirely my responsibility. The views expressed in this paper are the author's own and do not necessarily represent those of the institutions to which he is affiliated.

## 1. – Introduction

The optimal size of public sector is one of the most appealing topic in fiscal policy and public finance studies. Several theories have been advanced to explain this problem in different countries. Among them we find Wagner's Law (1912) of increasing state activities, Peacock and Wiseman hypothesis (1961), critical-limit hypothesis (Clark (1945; 1964)), Leviathan hypothesis (Brennan and Buchanan (1980)), differential productivity hypothesis (Baumol (1967)), and the relative price hypothesis (Balassa (1964); Samuelson (1964)). So, economic literature identified several determinants of public expenditure growth: inflation (Clark (1933; 1937; 1945; 1964)), total revenue (De Viti De Marco (1893; 1898; 1934); Dalena and Magazzino (2010)), debt service or burden ratio (Ricardo (1817); Barro (1974; 1989); Reinhart and Rogoff (2010)), GDP growth rate (Barro (1989; 1990); Scully (1994); Armey (1995); Forte and Magazzino (2010); Magazzino (2008; 2009b; 2010a; 2010b)), strategic transfers from federal government to the state governments, population growth, urbanization effect (Wagner (1912)), and taxation.

Over the past three decades, some studies – using the concepts of cointegration and Granger causality – focused on several countries and time periods. Yet, empirical findings are mixed and, for some countries, controversial. The results differ even on the direction of causality and the short-term versus long-term effects on economic policies. Depending upon what kind of causal relationship exists, its policy implications may be significant.

The aim of our study is to analyze the nexus between public expenditure and prices for the Mediterranean countries in the period 1970-2009. The data used in this work were taken from the IMF *Government Finance Statistics* database.

In addition, Italy has a high public debt to GDP ratio and a high share of public expenditure; so, the reduction in public expenditure could represent a valid way for the consolidation of public finances. However, reducing the size of public sector should focus on the expenditure items that have less impact on GDP.

The outline of this paper is as follows. Section 2 provides a survey of economic literature on this issue. Section 3 provides an overview of the applied empirical methodology and a brief discussion of the data used. Section 4 discusses the empirical results. Finally, Section 5 presents our concluding remarks.

## 2. – Literature Survey

Already Ricardo (1824) stressed the importance of separation of the central bank from political institutions, and the prohibition of monetary financing by the excesses of government spending (deficit monetization), only to clearly enunciate the key principles of the theory of today' central bank independence from political power.

The Australian statistical C. Clark (1933, 1937, 1945, 1964) warned the most economically advanced countries of the danger of letting go beyond the relationship between public expenditure and national income as the threshold value of 25%. Clark lays down that when government tax proceeds reach this critical ratio, a progressive tax system generates increasing proportions of additional income from taxpayers, whose productivity falls. In fact, high levels of taxation would have reduced incentives to work and saving. Moreover, people become less resistant to the inflationary methods of government financing.

According to the analysis of Clark, the higher taxes would have decreased the profits of enterprises, which, passing it on to prices, have increased the prices of final goods. So, the overall effect is a fall of the aggregate supply (due to the falls of private incentives) and an expansion of the aggregate demand (due to the inflationary financing techniques) and, hence, inflation results.

Through an analysis of time series on prices, taxes and public spending of a large group of countries for the inter-War period, Clark came precisely the threshold of 25% as a ratio of public expenditure on national income. If it is true that inflation is a "social evil", it is true that inflation reduces the costs of the public sector, since certain groups in society cannot defend. Moreover, the fiscal drag – the crop that inflation gives policy-makers in countries with progressive tax systems of type – is disappearing in many states, since the awareness of citizens in this respect has increased in recent years. Yet, recent decades have, however, proved that many countries have crossed the 25% limit without much inflationary tendencies (Jain (1989)).

Already Bernstein (1936) had investigated the possibility of using the public know which specific anti-recession tool, highlighting the effects of inflation. According to the scholar, in the first three decades of the twentieth century great attention was given to possible use of public expenditure in order to minimize cyclical fluctuations on employment and production (emphasized by the report of the "Royal Commission on the Poor Laws"), while others economists – as Keynes, Martin, Foster, Catchings and Pigou – had suggested the use of public spending as an instrument of economic policy, whereas periods of depression as a stage characterized by a low cost. Bernstein came to the conclusion that if these conditions were not favorable, considerable increases in public spending during periods of economic depression would lead to increased prices and production.

Bullock (1934), about the crisis of the thirties, put it on the rise-to the effects of economic policy choices of the Administration status, stressing the inadequacy and the lateness of the spending policies enacted in the years

1933-1934, also in view of the level of prices and sharp decline in tax revenue. Basically, if the start conditions of monetary stability are preserved, then the government will have ample room in the policies of deficit.

Pechman and Mayer (1952) discussed the limits to the inflation taxation outlined by Clark, concluding that in the period between the two world wars, the empirical evidence supports the thesis Clark in only two cases (Britain and Norway). Similarly, the price indices calculated for the period 1945-1948 grew annually in 53 of the 71 countries considered here: Clark arguments do not prove that prices grew faster where the tax burden exceeded the limit set by him.

Eltis (1983), analyzing the causes of the difficulties of the British economy in the seventies, found a double bond between inflation and public spending on the one hand, inflation was seen as the effect of deficit policies, useful - through increases supply of money - to finance the excess expenditure. Secondly it was originated by the wage increases put forward by workers to protect their purchasing power. Eltis found a strong empirical evidence to support the view that robust budget deficits inflationary pressures.

Tanzi, Blejer and Teijeiro (1987) moved from the consideration that the different parts of the public budget respond differently to inflationary pressures. However, scientists spotted in public debt service a strong link between public spending and the price trend.

Buiter (1987) studies the consequences for inflation of public expenditure cuts, emphasizing the important distinction between cuts in public consumption expenditure (which will tend to reduce the deficit) and cuts in public sector capital formation (which may have the perverse effect of increasing the deficit). This will happen if the expenditure effect is swamped by the direct and indirect effects of a reduced public sector capital stock on government revenues. If the public sector deficit increased, the cuts in public sector capital formation will raise the demand for seigniorage revenue.

Özatay (1997) studying the Turkish experience in the period 1997-1995 emphasizes the importance of coordination of fiscal and monetary policies in achieving price stability. Results indicate that, despite the rapidly changing financial environment, there are stationary long-run money-income relationships. Moreover, the growth rates of various monetary aggregates have predictive power for future movements in the Consumer Price Index. However, as the Turkish case clarifies, in an economy with persistent budget deficits these properties are not sufficient to conduct successful monetary policies. By a credible policy, it is possible to substantially reduce the inflation rate from 85% to 10% in a 4-year period. Yet, this necessitates that the Public Sector Borrowing Requirement should not exceed 1.5% of GNP.

Ruge-Murcia (1999) develops a dynamic, rational expectations model of inflation where the money supply is endogenously determined by the government's use of newly created money to finance its current spending and by the effect of past rates of inflation on the real value of taxes. In an empirical application to Brazil (1980-1989, monthly data), estimates indicate that there are steady-state inflation and money growth rates associated with each of the two possible government spending regimes. The low regime would be characterized in equilibrium by rates of inflation and money growth of 8.22% and 7.29% per month, respectively, and a share of GDP devoted to government outlays of 22.73%. The high spending regime would be associated with an expenditure level amounting to 33.43% of GDP, a monthly rate of inflation of 19.12%, and a monthly money growth rate of 19.25%.

Aizenman and Hausmann (2000), investigate budgetary rules for an economy characterized by inflation and volatile relative prices. In the absence of shocks, the design of the budget is that the Treasury allocates funds once in every budgetary cycle. In the presence of volatile shocks, one would observe occasional budgetary revisions, the outcome of which is that the actual expenditure differs from the projected one. They use a panel data for Argentina, Brazil, Chile, Columbia, Costa Rica, Caribbean, Salvador, Guatemala, Honduras, Mexico, Peru, and Venezuela, for 1970-1994. The correlation between the budget error and the inflation variable turned out to be high, and highly significant. Similar results are found for the case where inflation is decomposed into the expected and the unexpected components, confirming that both the expected and the unexpected inflation increase the budget error.

Alavirad (2003) studies the effect of inflation on government revenue and expenditure for Islamic Republic of Iran. His major finding is that the government budget deficit increases in the inflationary condition. In addition, the deficit increases money supply, and this tends to increase inflation in Iran.

Ezirim and Muoghalu (2006), starting from Clark's hypothesis, found that when the size of the public sector (measured by the share of expenditure on GDP) exceed a certain threshold, incentives to produce are discouraged (because of high tax burden). The reduction in aggregate supply, in addition, is even more pronounced in the case of budget balance (viewed as a fiscal constraint). The net result of such a bad adjustment between demand and supply is an inflationary spiral.

Kia (2006), studying Iranian economy for the period 1970-2002, focuses on internal and external factors, which influence the inflation rate in developing countries. According to the estimation results, over the long run, a higher exchange rate leads to a higher price in Iran. So, a policy regime that leads to a stronger currency can help to lower inflation. However, a higher money supply when it is anticipated does not lead to a higher price level, but an unanticipated shock in the money supply results in a permanent rise in the price level. So, an unanticipated reduction in the money supply should be a powerful tool to reduce inflation in Iran. It is also found that the fiscal policy is very effective in Iran to fight inflation as the increase in the real government expenditures as well as deficits cause inflation, but if the changes are unanticipated they cause the opposite effect. While a high debt per GDP is deflationary.

Ezirim, Muoghalu and Elike (2008) studied the relationship between growth rate of public spending and inflation rate for the United States of America in the period 1970-2002 found that the two variables move in the same direction. According to their analysis, inflation affects spending decisions of the U.S. federal government, but is in turn influenced both the short and long term. The dual causality was confirmed, however, the conclusions were reached and Ezirim and Ofurum (2003). The conclusion drawn by these scholars is that, in order to bend inflation, governments should appropriately reduce the levels of expenditure, on the other hand, to reduce the growth in the size of the public, policy-makers should diminish price dynamics. A further consequence would be that fiscal policy would be a valuable tool for controlling inflation, by virtue of their ability to act directly on public spending (content).

Pekarski (2010) analyzes budget deficits and inflation in high inflation economies. The main finding is that recurrent outbursts of extreme inflation in these economies can be explicitly explained by the hysteresis effect associated with the action of two mechanisms: the arithmetic of the wrong side of the ITLC and the Patinkin effect. Another finding is that changes in different items of the budget balance sheet may have very different effects on inflation (apart from their different effects on the real economy).

Varvarigos (2010) constructs a stochastic, dynamic general equilibrium model of endogenously sustained growth of an economy whose government finances volatile public spending via seigniorage. The resulting volatility in money supply, combined with the effects of money on human capital formation, yielded some interesting and important results concerning macroeconomic performance. The model predicts a negative correlation between long-run output growth and policy volatility. In addition, given that both the mean and the variance of the inflation rate are elevated by volatility in public spending, the model provides a possible account for the strong positive correlation between inflation and its variability, as well as their negative correlation with output growth.

## 3. – Data and methodology

For the purpose of this paper, the variables analyzed have been expressed in a logarithmic form. The data that have been used are annual and cover the time period 1970-2009, for Mediterranean countries.

The data used in this work were taken from the IMF *Government and Finance Statistics* database, which provide current and internationally comparable data on the finances and fiscal policies of Fund member governments[1].

Most of time series have unit root as many studies indicated, including Nelson and Plosser (1982), and as proved by Stock and Watson (1988) and Campbell and Perron (1991) among others, that most of the time series are non-stationary. The presence of a unit root in any time series means that the mean and variance are not independent of time. Conventional regression techniques based on non-stationary time series produce spurious regression and statistics may simply indicate only correlated trends rather than a true relationship (Granger and Newbold, 1974). Spurious regression can be detected in regression model by low Durbin-Watson statistics and relatively moderate $R^2$.

One of the most widely used unit root test is the ADF unit root test (Dickey and Fuller, 1979, 1981). Alternatively, Phillips (1987) and Phillips and Perron (1988) proposed a nonparametric method to correct a wide variety of serial correlation and heteroskedasticity (PP). Perron (1989, 1990) demonstrates that if a time series exhibits stationary fluctuations around a trend or a level containing a structural break, then unit root tests will erroneously conclude that there is a unit root. PP and ADF tests have the same asymptotic distributions.

Elliott, Rothenberg, and Stock (DF-GLS, 1996) proposed a modified Dickey-Fuller t test (known as the DF-GLS test). Essentially, the test is an augmented Dickey-Fuller test, except that the time series is transformed via a generalized least squares (GLS) regression before performing the test. The augmented Dickey-Fuller test involves fitting a regression of the form

$$\Delta y_t = \alpha + \beta y_{t-1} + \delta t + \xi_1 \Delta y_{t-1} + \xi_2 \Delta y_{t-2} + \ldots + \xi_k \Delta y_{t-k} + \varepsilon_t \qquad [1]$$

and then testing the null hypothesis $H_0$: $\beta$=0. The DF-GLS test is performed analogously but on GLS-detrended data. The null hypothesis of the test is that $y_t$ is a random walk, possibly with drift.

Finally, the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS, 1992) test differs from those unit root tests in common use (such as ADF, PP, and DF-GLS) by having a null hypothesis of stationarity. The test may be conducted under the null of either trend stationarity (the default) or level stationarity. Inference from this test is complementary to that derived from those based on the Dickey-Fuller distribution.

Then we examine the unit root (or stationarity) properties of the variables, accounting for structural breaks. The present paper employs Zivot and Andrews (ZA, 1992) test to address this issue. The Zivot and Andrews test is performed by running the following regressions:

$$x_t = \mu + \beta t + \alpha x_{t-1} + \sum_{i=1}^{k} \delta_i \, \Delta x_{t-i} + \varepsilon_t \qquad [2]$$

for $t$=1,…,$T$, where $x_t$ is a potentially non-stationary time-series, and the

[1] See: http://www.esds.ac.uk/international/support/user_guides/imf/gfs.asp.

terms $\Delta x_{t-i}$, $i$=1,…,$k$ are included to purge any serial correlation among residuals.

Furthermore, Clemente, Montañés and Reyes (CMR, 1998) have developed a procedure allowing for a gradual shift in the mean to test more than one break point.

The non-stationary series with the same order of integration may be cointegrated if there exist some linear combination of the series that can be tested for stationarity. The Johansen and Juselius procedure (Johansen, 1988; Johansen and Juselius, 1990) is preferable to test for cointegration for more than two series.

Moreover, Johansen and Juselius procedure is considered better than Engle-Granger even in two time series case and has better small sample properties since it allows feedback effects among the variables under investigation where it is assumed in the Engle and Granger procedure that there are no feedback effects between the variables. The procedure is based on likelihood ratio (LR) test to determine the number of cointegration vectors in the regression. Johansen technique enables to test for the existence of non-unique Cointegration relationships.

Three tests statistics are suggested to determine the number of cointegration vectors: the first is Johansen's "trace" statistic method, the second is his "maximum eigenvalue" statistic method, and the third method chooses $r$ to minimize an information criterion.

Having established the long-run equilibrium relationship between government expenditure and revenues, the short-run adjustments are estimated using the error correction model (ECM). The error correction model is based on the two following equations:

$$\Delta X_t = \alpha_0 + \alpha_1 e_{t-1} + \sum_{i=1}^{m} \alpha_i \Delta X_{t-i} + \sum_{i=1}^{n} \alpha_i \Delta Y_{t-i} + \varepsilon_t \quad [3]$$

$$\Delta Y_t = \beta_0 + \beta_1 u_{t-1} + \sum_{i=1}^{m} \beta_i \Delta Y_{t-i} + \sum_{i=1}^{n} \beta_i \Delta X_{t-i} + \eta_t \quad [4]$$

where $e_{t-1}$ and $u_{t-1}$ represent the error-correction terms which are the lagged residuals from the cointegration relations. The error correction terms will capture the speed of the short-run adjustments toward the long-run equilibrium. Furthermore, the error correction model equations (3) and (4) allow to test for short-run as well the long-run causality between government expenditure and aggregate income.

The short-run causality is based on a standard F-test statistics to test jointly the significance of the coefficients of the explanatory variable in their first differences. The long-run causality is based on a standard t-test. Negative and statistically significant values of the coefficients of the error correction terms indicate the existence of long-run causality.

## 4. – Econometric results

We present and discuss an empirical analysis of the nexus between public expenditure and inflation, applied to the Mediterranean countries. In Table 1 variables of the model are summed up. All series contains yearly data in real terms.

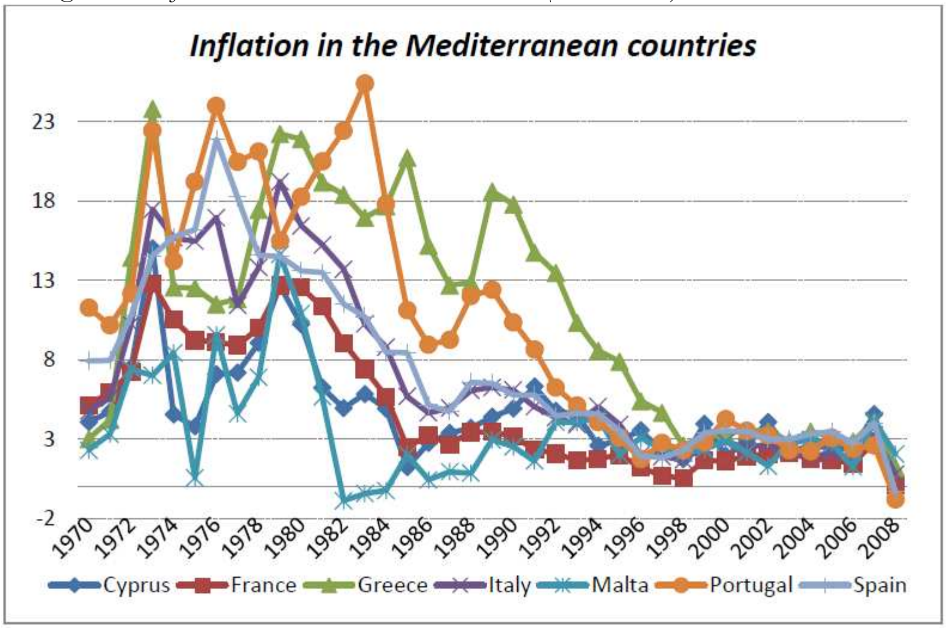
*Table 1 – List of variables.*

| *Variable* | *Explanation* |
|---|---|
| ***TEGG*** | Total Expenditure of General Government, % of GDP |
| ***NCPI*** | National Consumer Price Index, 2000=100 |

Source: IMF.

In Figure 1 the first differences of NCPI (*⊿NCPI*) for the Mediterranean countries from 1970 to 2009 are shown.

*Figure 1 – Inflation in the Mediterranean countries (1970-2009).*




Source: our elaborations on IMF data.

As a preliminary analysis, some descriptive statistics are shown in the following Table 2.

*Table 2 – Exploratory data analysis (mld EIT, 1970-2009).*

| *Variable* | *Mean* | *Median* | *Standard Deviation* | *Skewness* | *Kurtosis* | *Range* |
|---|---|---|---|---|---|---|
| ***TEGG*** | 45.2512 | 44.8566 | 5.5844 | -0.2273 | 2.5783 | 26.6596 |
| ***NCPI*** | 66.1517 | 71.2841 | 38.5341 | -0.1515 | 1.7357 | 130.3617 |

Source: our calculations on IMF data.

Correlation coefficients summarized in Table 3 indicate a low positive correlation between real total public expenditure and price index. These findings indicate that higher values of real public expenditure are associated with higher values of NCPI (except in Italy and Spain).

*Table 3 – Correlation between public expenditure and price index.*

| Country | Correlation coefficient between TEGG and NCPI | Correlation coefficient between ΔTEGG and ΔNCPI |
|---|---|---|
| ***Cyprus*** | 0.9095 | -0.0617 |
| ***France*** | 0.8344 | 0.3088 |
| ***Greece*** | 0.6284 | -0.1408 |
| ***Italy*** | 0.1241 | 0.3503 |
| ***Malta*** | 0.5106 | -0.0642 |
| ***Portugal*** | 0.9422 | -0.1384 |
| ***Spain*** | -0.1412 | -0.3202 |

Notes: Bonferroni adjustment applied.
Source: our calculations on IMF data.

Above all, we obtained log-transformations of the time-series. As a preliminary analysis, Inter-Quartile Range show the absence of outliers in our samples. Then, we applied time-series techniques on stationarity and unit root processes, in order to check some stationarity properties. Table 4 contains results of common unit root tests, for our variables.

*Table 4 – Results for stationarity tests.*

| Country | Variable | Stationarity tests | | | | |
|---|---|---|---|---|---|---|
| | | **Deterministic component** | **ADF** | **ERS** | **PP** | **KPSS** |
| ***Cyprus*** | ***TEGG*** | intercept, trend | NS: -2.357 | NS: -2.336 | NS: -2.357 | TS: 0.113 |
| | ***NCPI*** | intercept, trend | NS: -2.887 | NS: -1.217 | NS: -2.871 | NS: 0.357 |
| | ***ΔTEGG*** | intercept | DS: -3.418 | DS: -2.298 | DS: -3.418 | DS: 0.067 |
| | ***ΔNCPI*** | intercept | DS: -4.694 | NS: -1.929 | DS: -4.647 | NS: 0.534 |
| ***France*** | ***TEGG*** | intercept, trend | NS: -3.369 | NS: -2.465 | NS: -2.421 | NS: 0.188 |
| | ***NCPI*** | intercept, trend | NS: -2.128 | NS: -1.083 | NS: -1.388 | NS: 2.030 |
| | ***ΔTEGG*** | intercept | DS: -3.159 | DS: -2.631 | DS: -3.159 | DS: 0.129 |
| | ***ΔNCPI*** | intercept | NS: -1.942 | NS: -1.412 | DS: -1.923 | NS: 0.433 |
| ***Greece*** | ***TEGG*** | intercept | NS: -2.183 | NS: -0.247 | NS: -2.183 | NS: 0.555 |
| | ***NCPI*** | intercept, trend | NS: -2.619 | NS: -2.268 | NS: -2.676 | NS: 0.451 |
| | ***ΔTEGG*** | intercept | DS: -5.183 | NS: -2.059 | DS: -5.183 | DS: 0.138 |
| | ***ΔNCPI*** | intercept | NS: -1.610 | NS: -1.092 | NS: -1.579 | NS: 1.140 |
| ***Italy*** | ***TEGG*** | intercept | LS: -2.855 | NS: -0.733 | LS: -2.855 | LS: 0.269 |
| | ***NCPI*** | intercept, trend | NS: -2.153 | NS: -1.818 | NS: -2.937 | NS: 0.208 |
| | ***ΔTEGG*** | intercept | DS: -3.708 | NS: -1.481 | DS: -3.708 | DS: 0.352 |
| | ***ΔNCPI*** | intercept | NS: -2.622 | NS: -1.373 | NS: -2.567 | NS: 0.676 |
| ***Malta*** | ***TEGG*** | intercept | LS: -2.917 | NS: -1.360 | LS: -2.917 | LS: 0.399 |
| | ***NCPI*** | intercept, trend | NS: -2.399 | NS: -2.445 | NS: -1.549 | NS: 0.168 |
| | ***ΔTEGG*** | intercept | DS: -4.732 | NS: -1.937 | DS: -4.732 | DS: 0.131 |
| | ***ΔNCPI*** | intercept | DS: -3.715 | DS: -2.703 | DS: -3.725 | DS: 0.143 |
| ***Portugal*** | ***TEGG*** | intercept, trend | NS: -3.277 | TS: -3.434 | NS: -3.086 | NS: 0.148 |
| | ***NCPI*** | intercept, trend | NS: -2.120 | NS: -2.125 | NS: -2.972 | NS: 0.309 |
| | ***ΔTEGG*** | intercept | DS: -4.125 | DS: -2.783 | DS: -4.098 | DS: 0.064 |
| | ***ΔNCPI*** | intercept | NS: -1.527 | NS: -1.454 | NS: -1.640 | NS: 0.610 |
| ***Spain*** | ***TEGG*** | intercept | NS: 0.720 | NS: -1.389 | NS: -1.218 | LS: 0.232 |
| | ***NCPI*** | intercept, trend | TS: -4.910 | NS: -1.573 | TS: -4.711 | NS: 0.270 |
| | ***ΔTEGG*** | intercept | DS: -3.222 | DS: -2.321 | DS: -3.299 | DS: 0.458 |
| | ***ΔNCPI*** | intercept | NS: -2.575 | NS: -1.209 | NS: -2.697 | NS: 0.673 |

Notes: LS: Level Stationary; NS: Non Stationary; TS: Trend Stationary; DS: Difference Stationary.

Source: our calculations on IMF data.

The second column presents results for Augmented Dickey and Fuller (1979) test; the third one for Elliott, Rothenberg and Stock (1992) test; the fourth column contains results for Phillips and Perron (1988) test; at last, in the fifth column there are results for Kwiatkowski, Phillips, Schmidt and Shin (1992) test. Here, results indicate that public expenditure is clearly a *I(1)* process in five countries (Cyprus, France, Greece, Portugal and Spain); a *I(0)* process for Italy and Malta (where it seems to be level-stationary). While, prices index is a *I(2)* process everywhere, except Malta (*I(1)*).

*Table 5 – Results for unit root tests with structural breaks.*

| ***Country*** | ***Variable*** | ***TB*** | ***k*** | ***t-stat*** | ***1% Critical Value*** | ***5% Critical Value*** |
|---|---|---|---|---|---|---|
| ***Cyprus*** | ***TEGG*** | 2003 | 0 | -4.573 | -5.57 | -5.08 |
| | ***ΔTEGG*** | | 0 | -5.504 | -5.57 | -5.08 |
| | ***ΔNCPI*** | | 0 | -5.408 | -5.57 | -5.08 |
| | ***$\Delta^2$NCPI*** | | 1 | -6.428 | -4.93 | -4.42 |
| ***France*** | ***TEGG*** | 1992 | 1 | -3.424 | -5.57 | -5.08 |
| | ***ΔTEGG*** | | 0 | -5.752 | -5.57 | -5.08 |
| | ***ΔNCPI*** | | 2 | -3.918 | -5.57 | -5.08 |
| | ***$\Delta^2$NCPI*** | | 0 | -5.573 | -4.93 | -4.42 |
| ***Greece*** | ***TEGG*** | 2006 | 0 | -3.663 | -5.57 | -5.08 |
| | ***ΔTEGG*** | | 0 | -7.309 | -5.57 | -5.08 |
| | ***ΔNCPI*** | | 0 | -4.395 | -5.57 | -5.08 |
| | ***$\Delta^2$NCPI*** | | 0 | -5.601 | -5.57 | -5.08 |
| ***Italy*** | ***TEGG*** | 2005 | 0 | -2.508 | -5.57 | -5.08 |
| | ***ΔTEGG*** | | 0 | -6.282 | -5.57 | -5.08 |
| | ***ΔNCPI*** | | 0 | -4.679 | -5.57 | -5.08 |
| | ***$\Delta^2$NCPI*** | | 0 | -5.983 | -4.93 | -4.42 |
| ***Malta*** | ***TEGG*** | 2003 | 0 | -5.230 | -5.57 | -5.08 |
| | ***ΔTEGG*** | | 0 | -5.324 | -5.57 | -5.08 |
| | ***ΔNCPI*** | | 0 | -3.880 | -4.93 | -4.42 |
| | ***$\Delta^2$NCPI*** | | 2 | -4.892 | -4.93 | -4.42 |
| ***Portugal*** | ***TEGG*** | 1997 | 1 | -4.512 | -5.57 | -5.08 |
| | ***ΔTEGG*** | | 1 | -4.439 | -4.93 | -4.42 |
| | ***ΔNCPI*** | | 0 | -3.055 | -5.57 | -5.08 |
| | ***$\Delta^2$NCPI*** | | 2 | -4.835 | -4.93 | -4.42 |
| ***Spain*** | ***TEGG*** | 2007 | 0 | -1.008 | -5.57 | -5.08 |
| | ***ΔTEGG*** | | 0 | -7.911 | -5.57 | -5.08 |
| | ***ΔNCPI*** | | 1 | -2.892 | -5.57 | -5.08 |
| | ***$\Delta^2$NCPI*** | | 0 | -6.047 | -4.93 | -4.42 |

Source: our calculations on IMF data.

The results of the Zivot and Andrews unit root test are summarized in Table 5. An examination of these results for public expenditure series indicate that the null hypothesis of a unit root cannot be rejected in levels (the only exception is Malta, at a 5% significance level). If we take the first differences, we can reject the null hypothesis for all countries, unless for Spain. So, we can conclude that public expenditure is clearly a *I(1)* process in six countries (Cyprus, France, Greece, Italy, Portugal and Spain); a *I(0)* process for Malta. Inflation is a *I(1)* process everywhere.

*Table 6 – Results for additive outlier unit root tests.*

| ***Country*** | ***Variable*** | ***SB*** | ***k*** | ***t-stat*** | ***5% Critical Value*** |
|---|---|---|---|---|---|
| ***Cyprus*** | ***TEGG*** | 2000 | 0 | -3.366 | -3.560 |
| | ***ΔTEGG*** | | 0 | -6.378 | -3.560 |
| | ***ΔNCPI*** | | 4 | -3.637 | -3.560 |
| | ***$\Delta^2$NCPI*** | | 1 | -8.361 | -3.560 |
| ***France*** | ***TEGG*** | 1994 | 1 | -3.954 | -3.560 |
| | ***ΔTEGG*** | | 0 | -3.964 | -3.560 |
| | ***ΔNCPI*** | | 5 | -3.177 | -5.490 |
| | ***$\Delta^2$NCPI*** | | 1 | -4.237 | -3.560 |
| ***Greece*** | ***TEGG*** | 2008 | 0 | -4.184 | -3.560 |
| | ***ΔTEGG*** | | 0 | -6.796 | -3.560 |
| | ***ΔNCPI*** | | 5 | -2.300 | -5.490 |
| | ***$\Delta^2$NCPI*** | | 0 | -6.516 | -5.490 |
| ***Italy*** | ***TEGG*** | 1989, 1996 | 4 | -1.310 | -5.490 |
| | ***ΔTEGG*** | | 0 | -5.559 | -5.490 |
| | ***ΔNCPI*** | | 2 | -5.962 | -5.490 |
| | ***$\Delta^2$NCPI*** | | 2 | -3.591 | -3.560 |
| ***Malta*** | ***TEGG*** | 2000 | 2 | -3.472 | -3.560 |
| | ***ΔTEGG*** | | 1 | -3.891 | -3.560 |
| | ***ΔNCPI*** | | 0 | -4.627 | -5.490 |
| | ***$\Delta^2$NCPI*** | | 1 | -6.566 | -3.560 |
| ***Portugal*** | ***TEGG*** | 1987, 1992 | 0 | -2.330 | -5.490 |
| | ***ΔTEGG*** | | 1 | -5.001 | -3.560 |
| | ***ΔNCPI*** | | 5 | -3.127 | -5.490 |
| | ***$\Delta^2$NCPI*** | | 2 | -3.892 | -3.560 |
| ***Spain*** | ***TEGG*** | 1998, 2007 | 3 | -3.646 | -5.490 |
| | ***ΔTEGG*** | | 1 | -3.754 | -3.560 |
| | ***ΔNCPI*** | | 0 | -4.205 | -3.560 |
| | ***$\Delta^2$NCPI*** | | 2 | -3.658 | -3.560 |

Source: our calculations on IMF data.

From the Table 6 above, we note that the Clemente *at al.* test results are quite different to those found with the Zivot and Andrews test. For *TEGG*, despite the structural break, we are unable to reject the null hypothesis of a unit root in five countries (Cyprus, Italy, Malta, Portugal and Spain); as a conclusion, public expenditure seems to be a *I(1)* process in these countries, but a *I(0)* process in France and Greece. Inflation is *I(0)* for Cyprus, Italy and Spain, and *I(1)* otherwise.

The lag-order selection has been chosen according to the final prediction error (FPE), Akaike's information criterion (AIC), Schwarz's Bayesian information criterion (SBIC), and the Hannan and Quinn information criterion (HQIC).

Cointegration tests have been subsequently applied, in order to be able to find the long-run relationship between public expenditure variation (*ΔTEGG*) and inflation (*ΔNCPI*). As is shown in Table 7, Johansen and Juselius cointegration method suggests that there is one cointegrating relationship only in one case (Portugal). In fact, the trace statistic and the maximum-eigenvalue statistic reject *r*=0 in favour of *r*=1 at the 5% critical value only for this country. Yet, for Italy we have a contradictory result: in fact, the trace statistic suggests *r*=1, while the maximum-eigenvalue statistic suggests *r*=0. As in the lag-length selection problem, choosing the number of cointegrating equations that minimizes either the SBIC or the HQIC provides a consistent estimator of the number of cointegrating equations. Yet, all these criteria suggest a rank=0 for our data. While, for Portugal we find the presence of cointegration (rank=1).

*Table 7 – Results for cointegration tests between public expenditure growth and inflation (ΔTEGG and ΔNCPI).*

| ***Johansen and Juselius procedure*** | | | | |
|---|---|---|---|---|
| **Country** | **Trace statistic** | **Maximum-eigenvalue statistic** | **SBIC<br>HQIC<br>AIC** | **Rank** |
| ***Cyprus*** | 14.7664<br>(19.96) | 8.3949<br>(15.67) | 17.0911<br>16.8696<br>16.9294 | r=0 |
| ***France*** | 13.4491<br>(19.96) | 10.5467<br>(15.67) | 14.2574<br>14.1279<br>14.0688 | r=0 |
| ***Greece*** | 9.6624<br>(19.96) | 8.2397<br>(15.67) | 15.9234<br>15.7583<br>15.7246 | r=0 |
| ***Italy*** | 20.1376<br>(19.96) | 13.1048<br>(15.67) | 14.3923<br>14.2574<br>14.2003 | r=0 |
| ***Malta*** | 13.9444<br>(19.96) | 8.6749<br>(15.67) | 16.3768<br>16.1553<br>16.2152 | r=0 |
| ***Portugal*** | 2.5511<br>(9.42) | 2.5511<br>(9.24) | 15.1416<br>14.8875<br>14.7679 | r=1 |
| ***Spain*** | 10.0287<br>(19.96) | 7.5472<br>(15.67) | 14.8161<br>14.5946<br>14.6545 | r=0 |

Notes: 5% Critical Values in parenthesis.
Source: our calculations on IMF data.

Granger causality tests suggest a bi-directional flow, at 1% significance level, for public expenditure growth and inflation in the Italian case, in the short-run; a unidirectional flow, in the direction from inflation to public expenditure for Portugal (in the long-run), and for France in the short-run (at 1% level); a unidirectional flow, but in the opposite direction (from public expenditure to inflation), for Cyprus (at 10% level), Malta (at 1% level) and Spain (at 5% level, see Table 8).

*Table 8 – Results for short and long-run causality tests.*

| ***Country*** | ***Lags*** | ***Log-likelihood*** | ***SBIC*** | ***Causality in the long-run*** | ***Causality in the short-run*** |
|---|---|---|---|---|---|
| ***Cyprus*** | 2 | -84.4065 | 16.5527 | - | ΔG → ΔP |
| ***France*** | 1 | -201.1204 | 14.3150 | - | ΔP → ΔG |
| ***Greece*** | 1 | -148.5432 | 16.0526 | - | - |
| ***Italy*** | 2 | -175.1035 | 14.4355 | - | ΔG → ΔP<br>ΔP → ΔG |
| ***Malta*** | 2 | -80.4865 | 15.8993 | - | ΔG → ΔP |
| ***Portugal*** | 1 | -221.5129 | 15.1773 | ΔP → ΔG | - |
| ***Spain*** | 1 | -82.4877 | 14.2689 | - | ΔG → ΔP |

Source: our calculations on IMF data.

For all our equations, a Lagrange-multiplier (LM) test for autocorrelation in the residuals of Vector Error-Correction Model (VECM) clarifies as at the 5% significance level we cannot reject the null hypothesis that there is no serial correlation in the residuals for the orders 1,…,5 tested. Checking the eigenvalue stability condition in a VECM, the eigenvalues of the companion matrix lie inside the unit circle, and the real roots are far from 1. As regard the Wald lag-exclusion statistics, we strongly reject the hypothesis that the coefficients either on the first lag or on the second lag of the endogenous variables are zero in all two equations jointly. The Jarque and Bera normality test results present statistics for each equation and for all equations jointly against the null hypothesis of normality. For our models, results suggest normality. Finally, the analysis of ARCH effects shows the absence of this problem for the estimated models.

## 5. – Conclusions and policy implications

The purpose of this paper is to contribute to the literature on the nexus between public expenditure and inflation, using recent econometric techniques. So, we studied the relationship between public expenditure and inflation for Mediterranean countries, using annual data covering the period 1970-2009. The time-series properties of the data were assessed using several unit root tests (ADF, DF-GLS, PP, and KPSS). Furthermore, in order to evaluate the presence of eventual structural breaks, some tests (ZA and CMR) have been conducted. Empirical findings indicate that public expenditure is clearly a *I(1)* process in five countries (Cyprus, France, Greece, Portugal and Spain); and a *I(0)* process for Italy and Malta. While, prices index is a *I(2)* process everywhere, except Malta.

Cointegration analysis reveals that there is a long-run relationship between public expenditure/GDP growth and inflation only for Portugal. Moreover, Granger causality tests suggest a bi-directional flow, at 1% significance level, for public expenditure growth and inflation in the Italian case, in the short-run; a unidirectional flow, in the direction from inflation to public expenditure for Portugal (in the long-run), and for France in the short-run (at 1% level); a unidirectional flow, but in the opposite direction (from public expenditure to inflation), for Cyprus (at 10% level), Malta (at 1% level) and Spain (at 5% level).

Yet, we find no clear evidence of government spending causing prices dynamics as far as the vice versa. In other words, the original Clark's proposition of an excessive government spending as a cause of pressure on prices in the economy is not completely supported by the data for Mediterranean countries. Certainly, this result is subject to the time period examined and statistical methods used; nevertheless, our empirical findings don't show a clear evidence in favour of the opposite direction of causality flow. In fact, inflation Granger-cause public expenditure growth in few cases.

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