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DOES THE FISHER EFFECT APPLY IN GREECE? A COINTEGRATION ANALYSIS*

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Abstract

In this paper we tested the joint hypothesis of the Fisher effect and rationality of inflation expectations in Greece during the period 1980:I - 1996:II applying cointegration technique.

The basic evidence of this paper is the invalidity of the Fisher relationship as a long-run equilibrium phenomenon in the case of the Greek Economy. This means that the nominal interest rate does not follow the interest rate changes over the long-run. Inflationary movements have not been totally absorbed by nominal interest rates and as consequence the Fisher effect is not valid. This failure implies that external factors play a direct role in the determination of the domestic interest rate, something which is reasonable for an open economy, such as the Greek economy, where capital flows are not prohibited. (JEL: C32)

1. Introduction

Much empirical work has been done on testing the Fisher hypothesis. Fisher's hypothesis continues to have a vast empirical interest in the literature. The Fisher hypothesis is the foundation of any economic theory on interest rates. Fisher (1930) argued that over the long-run, changes in inflationary expectations are perfectly absorbed by nominal interest rate under the condition of efficient capital markets. The implication of the validity of the Fisher hypothesis is that the real interest rate is determined only by real factors and cannot be influenced by monetary policy. Therefore, if expected inflation implies real interest rates, the Fisher hypothesis (Fisher Effect) is invalid.

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The empirical evidence, however, of early works failed to support a one to one relationship between inflation and nominal interest rate (Fama, 1975, 1977, Cargil, 1976, Carlson, 1977, Lahiri, 1979, Levi and Makin, 1978, Tanzi, 1980, Fama and Schwert, 1977). This failure of the Fisher hypothesis led Carmichael and Stebbing (1983) to support an alternative to the Fisher hypothesis which is known as the inverted Fisher hypothesis (IFH). According to the model of Carmichael and Stebbing, the nominal interest rate on financial assets is constant, with the real interest rate moving inversely one for one with the expected inflation and therefore the Fisher effect does not hold. Gallagher (1986) was the pioneer who attempted to clarify this conclusion. However the serious disandvantage of all the early empirical works was that they did not discriminate between the short and long-run adjustment process, and in addition do not cope with the misspecification problems arising from the ignorance of the statistical properties of the variables, and of the cointegration tests of inflation and nominal yields. Furthermore theoretically, the Fisher hypothesis was regarded as a long-run equilibrium phenomenon (Summers, 1983). The question of whether the Fisher hypothesis is valid should be answered by testing whether the nominal interest rate moves together with inflation over the long-run. For this reason many approaches have been applied for testing the long-run relationship between nominal interest rates and expected inflation (Lucas, 1980, Summers, 1983, Lothian, 1985).

However, only recently cointegration analysis for testing of long-run relationships between non-stationary variables was used to test the Fisher hypothesis (Atkins, 1989, MacDonald and Murphy, 1989, Moazzami, 1991, Gupta and Moazzami, 1990, Inder and Silvapulle, 1993, Garcia and Zapata, 1991, Bonham, 1991, Thornton, 1996, Daniels et al. 1996). The majority of these works support the validity of a long-run relationship between the nominal interest rate and inflation.

The scope of this paper is to examine if the Fisher effect holds for the Greek economy, which amounts to test the joint hypothesis of the Fisher hypothesis and rationality of inflation expectations in Greece, using cointegration techniques. The remainder of the paper is organised as follows: section 2 briefly presents the model specification of the Fisher relationship in order to test it empirically, as well as the data used for the estimation of the model. In section 3, we investigate the time series statistical properties of the data by employing recent developments in the econometrics of non-stationarities. Unit root tests that permit both stationarity and non — stationarity about either mean or trend are used following the approaches of Dickey and Fuller (Fuller, 1976, Dickey and Fuller, 1979, 1981), and Phillips and Perron (Perron, 1988, Phillips and Perron, 1988, Phillips, 1986). In addition, a test for the existence of structural breaks is used (Banerjee et al. 1992), because as Sargent (1976) supports, the presence of structural breaks in the nominal interest rate-inflation relation affects the relationship between interest and inflation rates[1]. This means that changes in monetary policy and exchange rate regimes might affect the Fisher relationship. In the next section, we apply the maximum likelihood cointegration technique proposed by Johansen (1988, 1991) and Johansen and Juselius (1990), to test for the possibility of cointegrating vector (a common stochastic trend) between the nominal interest rate and inflation and also we present the empirical results. Finally, the last section will be present the main conclusions of this paper.

2. Model Specification

According to Fisher (1930), if the market is efficient the expected nominal interest rate is equal to the sum of the expected real interest rate and expected rate of inflation. This proposition is written as:

$$i^e_t = r^e_t + \pi^e_t \tag{1}$$

where it is the nominal interest rate, at time t, r_t is the real interest rate at time t, π_t is the rate of inflation at time t and e is an expectations operator. According to the Fisher hypothesis, in long-run equilibrium a change in the rate of growth of money supply leads to a fully perceived change in inflation and an adjustment of nominal interest rates, with the consequence that the real interest rate will remain constant in the long-run (i.e. it is generated by a stationary process). Moreover, for nominal assets, nominal interest rates are contracted in advance so that the ex post nominal yield is known (Fama, 1975), and therefore is $i^e_t = i_t$.

The Fisher hypothesis specifies the following regression equation:

$$i_t = \alpha + \beta\pi^e_t + u_t \tag{2}$$

where u_t is an error term. Equation (2) is based on the assumption that ex ante real interest rate is generated by a stationary process (real expected rate of return is constant but subject to random error, $r^e_t = r_t + v_t$, $v_t \sim N(0, \sigma^2_v)$. The error term u_t in equation (2) includes the zero-mean shock to the ex ante real interest rate in period t. Given that the Fisher hypothesis is a long-run equilibrium phenomenon, coefficient β is the long-run effect of π^e on i. The Fisher hypothesis supports that β=1 in equation (2) which means that a long-run unit proportional equilibrium relationship exists between π^e and i.

Recent empirical analysis of the Fisher relationship usually supposes that expectations are rational (i.e. $\pi^e_t, = E(\pi^e_t / I_t)$), where E shows the conditional expectations operator and I_t is an information set available when expectations are formed. The implication of the rational expectations is that $\pi_t = \pi^e_t + \varepsilon_t$, where ε_t is stationary white noise error that captures all non-systematic measurement errors, (Rose, 1988). Equation (3) represents the form of the cointegrating regression which was finally estimated.

$$i_t = \alpha + \beta\pi_t + z_t \quad (3)$$

The error term z_t in equation (3) is the sum of stationary components (the inflation expectation error and the zero-mean shock to the expected real interest rate in period t).

In this paper, we will try testing equation (3) as a two-equation vector autoregressive (VAR) system using a two-equations VAR approach as advanced by Johansen (1988, 1991) and Johansen and Juselius (1990).

In this study, we investigate the period 1980: Q_1 to 1996: Q_2 using quarterly data. We used the three-month Treasury bill as the nominal rate of interest (i), which was not taxed until 1st of January 1997; the consumer price index (CPI) was used to construct the rate of inflation (π). Both variables were taken from the Monthly Statistical Bulletin of the Bank of Greece (various issues).

The reason that led us to use lower frequency data in the present study is that according to Shiller and Perron (1985) and Lothian and Taylor (1992) what appears to have importance is the total length of the sample period when investigating the long-run statistical properties of time series.

Computer packages which we have used are RATS 4.20 (1995), CATS in RATS (1995) and Microfit 3.86 (1992).

3. Univariate Properties of the Used Data

Model (3) is essentially a long-run equilibrium relationship. If the above equilibrium model holds, the variables included in the model must be cointegrated even if the individual variables are non-stationary (Engle and Granger, 1987).

However, before proceeding to investigate the existence of cointegrated vector in the next section, it is necessary to establish the time series properties of the individual series used by means of Dickey-Fuller (DF), Augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1979, 1981) and Phillips-Perron (PP) tests (Perron, 1988, Phillips, 1987, Phillips and Perron, 1988). Also we apply the test for the existence of structural breaks in the data, using the approach advanced by Banergee et al. (1992)[2].

The DF and ADF statistics for unit roots rely on a parametric approach. However, since this method reduces the efficiency of the test (DeJong et al, 1992), recently Phillips and Perron (Phillips, 1987, Perron 1988, Phillips and Perron, 1988) have advanced non-parametric tests for serial correlation and heteroscedasticity. Nevertheless we decided to present the Dickey - Fuller results in addition to the Phillips - Perron tests because the Dickey - Fuller and Augmented Dickey - Fuller statistics have better small-sample properties (Campbell and Perron, 1992). Dickey-Fuller and Augmented Dickey-Fuller tests are calculated by running the following regression:

$$\Delta y_t = \beta_0 + \beta_1 y_{y-1} + \beta_2 t + \sum_{i=1}^{P} \gamma_j \Delta y_{t-j} + \upsilon_t, \qquad \upsilon_t \sim N(0, \sigma^2) \tag{4}$$

where Δ denotes the first differences of the variables, t is a time trend and υ is a stochastic term, and testing the significance of β_1, by comparing the t — value of coefficient β_1 to the critical value of τ, in Fuller (1976, p. 373). If we accept the null hypothesis that $\beta_1 = 0$ in equation (4) then this means that Δy_t is stationary implying that $y_t \sim I(1)$ which means that y_t is non-stationary, while if we accept the alternative hypothesis that $\beta_1 < 0$ then this means that the hypothesis of a unit root for the series Δy_t is accepted. The number of lags, p, was chosen on the basis of being sufficiently large, so that the error term υ_t to be white noise, however taking into account to use only lags that were statistically significant[3].

The Phillips-Perron (PP) test involves computing the OLS regression[4].

$$y, = \mu + \alpha y_{t-1} + \beta (T - t/2) + \upsilon, \tag{5}$$

where υ is allowed to follow a wide variety of stochastic behaviour.

The null hypothesis, $\alpha = 1$, on (5) is tested by the statistic $Z(\tau_t)$ which involves a long algebraic expression (Baillie and McMahon, 1992, pp. 148-9) and has the Dickey - Fuller (τ_t) distribution. The null hypothesis is that the variable in question is stationary in first differences, $\Delta y_t \sim I(0)$. Table 1 shows the unit root tests.

Table 1 shows that the variables i and π, which are included in the Fisher's relationship, equation (3), are stationary in first differences (i.e. $\Delta i \sim I(0)$, $\Delta\pi \sim I(0)$). According to Nelson and Plosser (1982), most macroeconomic variables have a unit root (a stochastic trend), a concept which was challenged by Perron (1989). Perron supported that only certain «big shocks» have permanent effects on the various macroeconomic variables and that these shocks are exogenous. However, modelling such shocks as points of structural change in the economy generally result to a rejection of the null hypothesis of a unit root. The disadvantage of Perron's approach is the assumption that the data of the structural break is known a priori or that the data chosen is uncorrelated with the data. This was challenged by Christiano (1988), Zivot and Andrews (1992) and Banerjee et al. (1992) who, unlike Perron, treat the choice of the break point to be endogenous, which circumvents the problem of data-mining[5].

In addition to the standard ADF unit roots test, we used the technique of Banergee et al. (1992) in order to examine the stochastic properties of the variables by investigating the existence of structural breaks in the time series included in equation (3). The test of Banerjee et al. consists of comparing the results from a sequence of recursive ADF statistics including a constant and a time trend.

Table 2 presents the results on unit roots based on recussive test statistics.

Since the values reported in the table 2 are lower, in absolute terms, than the critical ones, then there is evidence that the acceptance of the null hypothesis of non — stationarity in the full sample should not be attributed to the presence of structural breaks, a fact which is consistent with the results of Table 1.

4. Johansen's Cointegration Test - Empirical Results

The basic idea behind cointegration is that if, in the long-run, two or more variables move closely together, the linear combination between them is stationary and hence we may consider those series as defining a long-run equilibrium relationship. Johansen starts by defining an η-dimensional vector of I(1) variables X. In our case this vector includes the variables in model (3). The vector autoregressive (VAR) representation of such a system with Gaussian errors ε_t is given as follows:

$$X_t - \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \ldots + \Pi_\varkappa X_{\nu-\varkappa} + \mu + \Theta D + \varepsilon_t \qquad (6)$$

where the Π_j are nxn coefficient matrices, ε_t is an independently **and identically** distributed η-dimensional vector with zero mean and covariance matrix Ω, μ is an nxl vector of unrestricted constant terms, (Johansen, 1992) and **D** is **centered** seasonal dummies which together with the constant terms account for short-run effects. The long-run coefficient matrix Π (stochastic cointegrating matrix) corresponding to (6) is defined by

$$\Pi = 1 - \Pi_1 - \Pi_2 \ . \ . \ . - \Pi_\varkappa \qquad (7)$$

where I is the identity matrix.

The system (6) can be reparameterised to yield a generalised error correction model:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \ . \ . \ . + \Gamma_{k-1} \Delta X_{t-k+1} + \Gamma_k X_{t-k} + \mu + \Theta\Delta + \ \varepsilon_t \quad (8)$$

where $\Gamma_i = -(I - \Pi_1 - \Pi_2 - \ldots - \Pi_i)$, $i = 1, 2, \ldots \varkappa-1$ and $\Gamma_k = -\Pi$ (Rank of matrix Π). The matrix of coefficients, Γ_k, defines the long-run solution to model (6). In essence, Johansen's cointegration test involves determining the rank of matrix Π (denoted r) in order to test the hypothesis that there is at least one cointegrating vector in the process governing movements of i_t and π_t.

There are three possibilities of the rank of Π:

(i) Rank (Π) = 0. The matrix Π is the null matrix which means that the variables X are not cointegrated, which implies that there is no long-run relationship between them.

(ii) Rank (Π) = n, (full rank), which means that the vector X is stationary, something which would contradict our earlier results that inflation and nominal interest rate are both integrated of order 1 (in our case is η =2).

(iii) Rank (Π) = r<n. The matrix Π has not full rank, implying that there are r statistically significant cointegrating vectors. (In our case will be r=1).

The matrix Π can be regarded as the product of two matrices (nxr), α and β as follows:

$$\Pi = \alpha\beta' \qquad (9)$$

where β contains the cointegrating vector and α is made up of the corresponding error-correction coefficients.

Johansen (1988, 1991) and Johansen and Juselius (1990, 1992) have constructed a maximum likelihood approach for estimating α and β, as well as a likelihood ratio (LR) test for detecting the number of significant cointegrating vectors (r). There are two statistics from the Johansen vector autoregressive tests that determine the rank of the cointegration space. One is the value of the LR test based on the maximum eigenvalue (λ_{max}) of the stochastic matrix. The other is the value of the LR test based on the trace of the stochastic matrix (Trace). The LR test statistic developed by Johansen for the hypothesis that there are at most r cointegrating vectors is as follows:

$$LR_{Tr} = -2\log(Q) = -T \sum_{i=r+1}^{n} \log(1 - \lambda_t), \qquad \text{(Trace Test)} \qquad (10)$$

where $\lambda_{r+1}, \lambda_{r+2}, \ldots, \lambda_n$ are the n-r smallest eigenvalues.

Johansen also considers the following LR test statistic for the hypothesis that there are r cointegrating vectors against the alternative of r+1:

$$LR_{max} = -2\log(Q) = -T\log(1 - \lambda_{r+1}), \quad \text{(Maximum Eigenvalue Test)} \quad (11)$$

In our case, the null hypothesis under the trace test is that the matrix Π is of rank 0, (r=0), and the alternative hypothesis is that r 1. The maximal eigenvalue test, (λ_{max}), is similar to the trace test, except that the alternative hypothesis would be that r= 1.

However, in order to apply the Johansen technique, a lag length must be selected for the VAR model. The approach used in specifying the number of lags in ΔX in equation (8) was based on Sims' (1980) likelihood ratio (LR) statistic taking into account the degrees of freedom (d.f.). We also used the Ljung-Box Q-statistic for serial correlation in the equations of the VAR model.

We initially decided to start using a model with eight lags ($\varkappa$= 8), under which the equations of the model are free from serial correlation. In addition a model with a six lag structure was estimated and tested against the model with eight lag structure. The choice of the lag structure was based exactly on these two selection criteria.

Table 3 presents the residuals misspecification tests and test of the lag structure for the VAR model.

Table 3 shows that the marginal level of significance for both of the statistics (Ljung-Box statistic and LR Sims statistic) is higher than 5% for the model with six lags structure. We therefore decided to use this lag structure, $\varkappa$=6 in our VAR model. From this table it is obvious that $\varkappa$=6 gave VAR equations which satisfied a range of diagnostic tests for misspecification. For the estimation of the model we have included three centered seasonal dummies while a constant is allowed to enter unrestricted in the VAR model[6].

The results from the trace and maximal eigenvalue tests (Johansen's cointegration tests) are presented in Table 4.

From the results of Table 4, we accept the null hypothesis (Ho) of no common trend (i.e. the matrix Π is of rank 0, r=0) at the 5% level in both the trace and maximal eigenvalue tests. The results failed to accept the hypothesis that there is a unique cointegrating vector relating the nominal rate of interest and inflation. This implies that the matrices β and α which contain the cointegrating vector and the corresponding error-correction coefficients, respectively, can not be found. Accordingly, we conclude that the nominal rate of interest and the inflation rate are not cointegrated, and as consequence there can be no error correction representation within a dynamic model showing the short- and long-run adjustment processes. (Alogoskoufis and Smith, 1991, Cuthbertson et al., 1992). Also, we are not in a position to identify the direction of causality, in Granger sense, between the nominal interest rate and the inflation rate (Engle and Granger, 1987).

5. Conclusions

The paper tested the joint hypothesis of the Fisher hypothesis and rationality of inflation expectations in Greece for the period 1980: Ql - 1996: Q2 using cointegration methodology.

The main conclusion of this paper is that the Fisher relationship could not be regarded as a long-run equilibrium phenomenon in the case of the Greek economy, which means that the nominal interest rate does not move together with the inflation rate over the long-run. Inflationary movements have not been totally absorbed in nominal interest rates and the Fisher effect does not hold.

This failure implies that external factors have a direct role to play in the determination of the domestic interest rate, something which is reasonable for an open economy, such as the Greek economy, where capital flows are not prohibited. An indirect implication of the above mentioned findings is that the power of monetary policy to affect the nominal interest rate, through inflation, is limited.

In addition, we point out that the implementation of economic policy became more difficult mainly during recent years due to the following reasons:(a) The financial liberalisation which started in the mid-1980's, and (b) the exchange rate policy (policy of "hard" drachma) replaced the traditional intermediate monetary target (i.e. money supply, M3). Figures 1 and 2 show the application of the liberalised approach on interest rates as well as the strict adherence to the "hard" drachma policy.

TABLE 1

Testing for Unit Roots: 1980: Q_1 - 1996: Q_2

	Dickey - Fuller (DF)		Augmented Dickey-Fuller (ADF)		Lagrange Multiplier (LM)	Phillips - Perron (PP)	
Variables	$DF(\hat{\tau}_\mu)$	$DF(\hat{\tau}_\tau)$	$DF(\hat{\tau}_\mu)$	$DF(\hat{\tau}_\tau)$	LM(4)	$Z(\hat{\tau}_\mu)$	$Z(\hat{\tau}_\tau)$
i	-1.2911(0)	-0.4945(0)	-	-	0.865	-1.2783	-2.5998
π	-	-	-1.7029(5)	-3.3165(5)	0.560	-1.4932	-0.4055
Δi	-7.9335(0)	-8.3680(0)	-	-	0.966	-6.4200	-6.4062
$\Delta\pi$	-6.3680(0)	-6.3041(0)	-	-	0.370	-8.0646	-8.0597

Figures in parentheses show the number of lagged dependent variables in the regression. The choice between zero and non-zero lags was based on the Lagrange multiplier (LM) test fourth-order serial correlation of the residual. The LM statistic is asymptotically distributed as $X^2(4)$ (d.f.= 4).

The numbers in the columns headed Phillips-Perron are the Dickey-Fuller statistic with the transformation suggested by Phillips (1987).

(τ_μ) and (τ_τ) are the test statistics allowing for constant mean, and for a time trend in mean, respectively. Approximate 5% critical value for (τ_μ) and Z(τ_μ) is -2.89 for a sample size of n=100, and the 5% critical value for (τ_τ) and Z(τ_τ) is -3.43 (Fuller, 1976, p. 373).

The calculation of PP statistics Z(τ_μ) and Z(τ_τ) were based on 4 time lags.

Figures in the column LM (4) show the marginal levels of significance.

TABLE 2

Results of Unit Roots Based on Recursive Test Statistics

Variables	t_{min}	t_{max}
i	-2.4747 (1993: Q_4)	-0.11667 (1989: Q_1)
π	-3.4534 (1995: Q_1)	0.80217 (1986: Q_1)

The estimates correspond to the maximum and minimum ADF t statistics values resulting from estimation starting at 1983: Q_4 recursively, using 5 lags in the ADF statistics. The dates at which these results were taken are reported in parentheses. Critical values at the 5% level are t_{max} = -1.99, and t_{min} = -4.33 (Banerjee et al. 1992, Table 1).

TABLE 3

Test of the Lag Structure and Residuals Misspecification Tests on the VAR Model (Equation 8)

Dependent Variable	SEE	Q(21)	N(2)	H(1)	σ^2	LR Sims Test
i	1.4329	20.4420	4.8831	0.0874	2.0532	7.4471
		(0.4934)	(0.102)	(0.767)		(0.4892)
p	0.8490	9.0058	3.7802	0.0513	0.7208	Lags with restri-
		(0.9891)	(0.151)	(0.821)		ctions/ Lags with-
						out restrictions 6/8

SEE is the standard error of the equation.

Q(21) is the Ljung-Box statistic for serial correlation with 21 d.f. It is distributed as $X^2(21)$.

N(2) is the Bera and Jarque (1980) statistic (BJ) for normality of the error terms. It is distributed as $X^2(2)$ with 2 d.f.

H(1) is the Lagrange multiplier statistic for the heteroscedasticity among the residuals. It is distributed as $X^2(1)$ with 1 d.f.

σ^2 shows the variance of the equation.

Numbers in parentheses are the marginal levels of significance.

LR in the last column is the likelihood ratio Sims test for the selection of the lag structure of a VAR model. It is distributed as $X^2(8)$ with 8 d.f., where the d.f. are equal to the number of restrictions.

TABLE 4

Johansen Cointegration Tests between i, π

Ho	n-r	Tr	95%	Ho		λ_{max}	95%
$r \leq 1$	1	2.5622	3.7620	$r \leq 1$	$r \leq 2$	2.5622	3.7620
$r=0$	2	12.1970	15.4100	$r=0$	$r=1$	9.6348	14.0690

r and (n-r) indicate the number of eigenvectors and common trends, respectively.

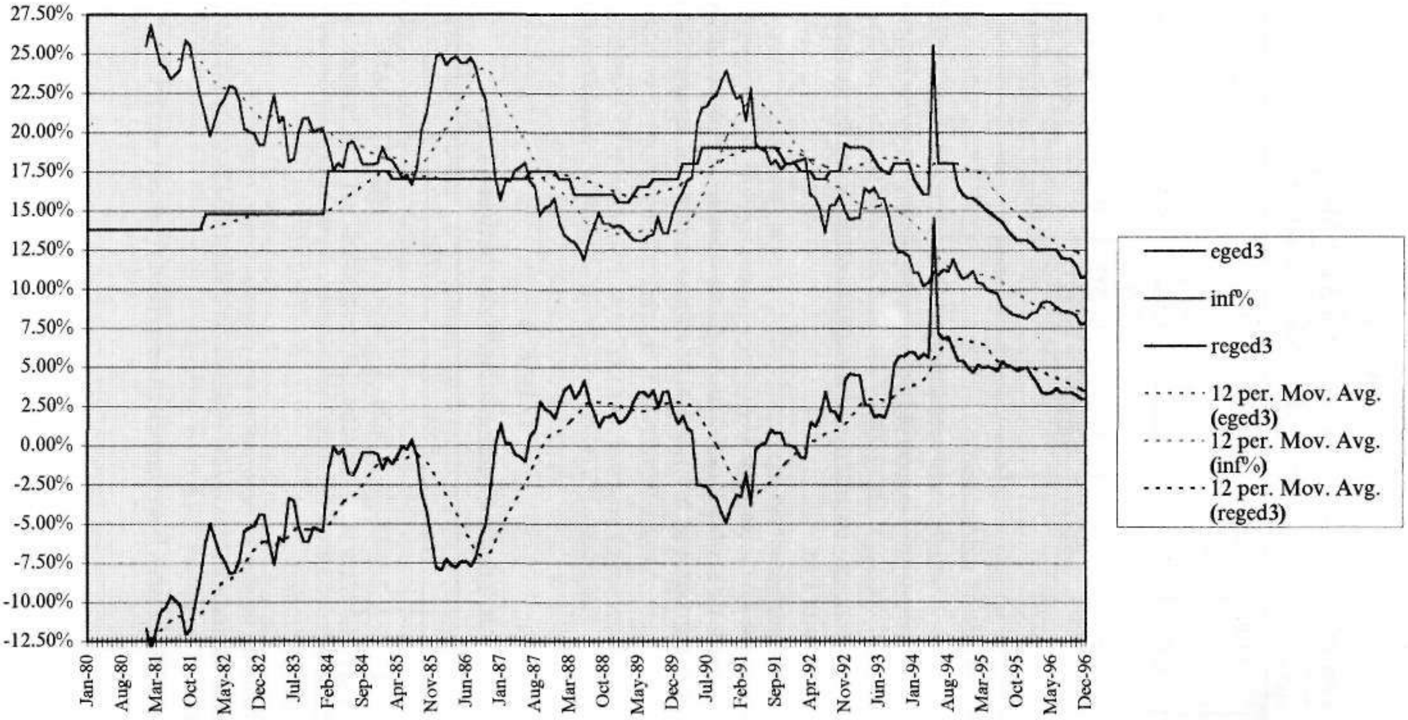
Tr and λ_{max} show the trace and maximum eigenvectors statistics respectively for the unrestricted model.

Critical value at 95% are taken from Osterwald-Lemun (1992) (Tables 1 and 1*).

Figure 1

o

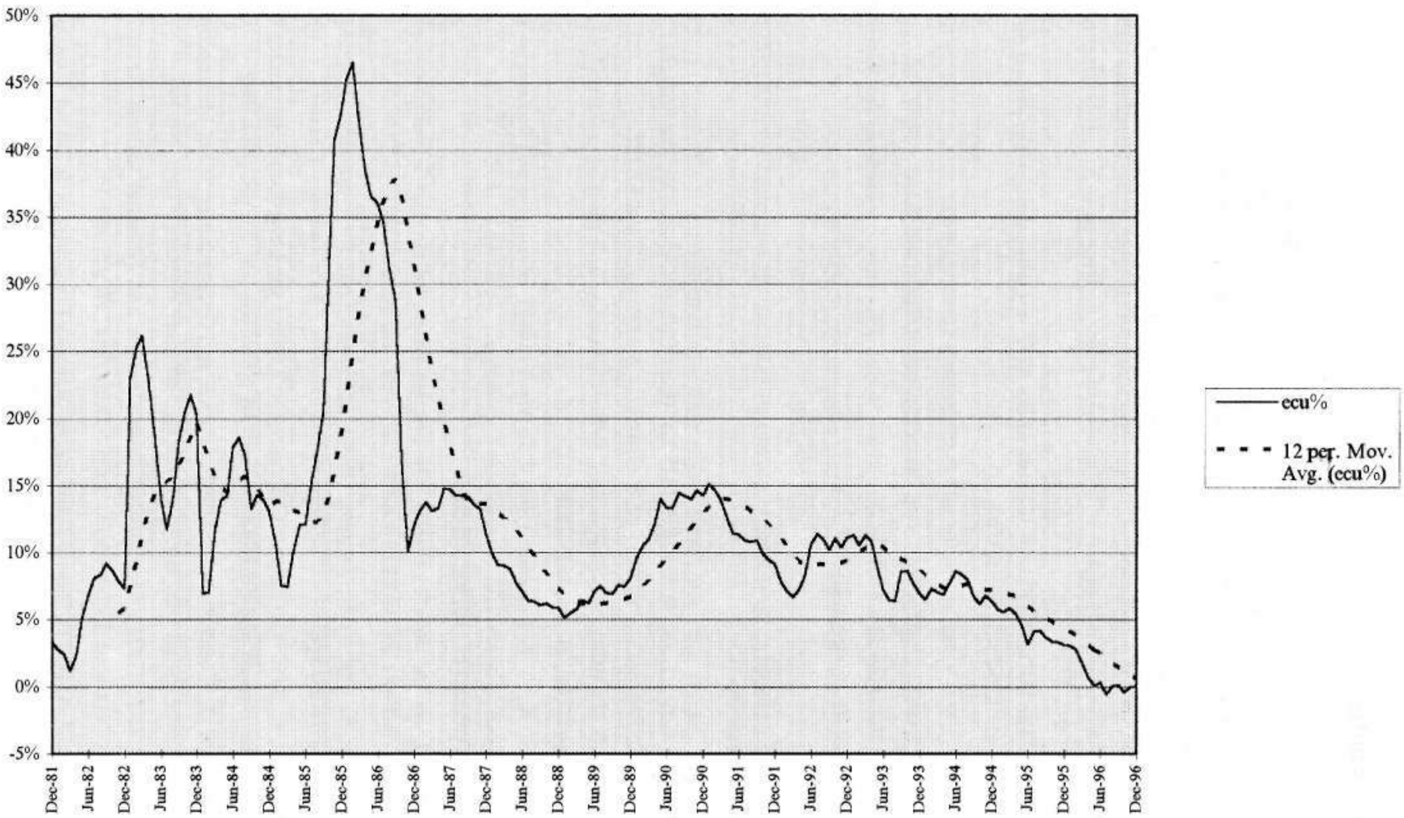
3-Month T-Bill - Inflation - Real Interest Rate



eged3:3-month TB inf%: inflation reged3: real interest rate

Figure 2

Drachma Depreciation Against ECU



Footnotes

1. See also Barth and Bradley, (1988), Hutchison and Keeley, (1989).

2. A stochastic process is said to be stationary if the means and variances of the process are constant over time while the auto covariance between two periods depends only on the gap between those periods and not the actual time at which the covariance is considered. Intuitively, if a time series, defined as a single realisation of a stochastic process, has a tendency to return to a fixed value over time, then it is stationary. If a non-stationary series X has to be differenced d times in order to induce stationarity it is said to be integrated of order d, denoted as $X_t \sim I(d)$ (Granger, 1983, Engle and Granger, 1987).

3. The DF test is a restricted case of the ADF test when the number of first differenced terms on the right hand side of equation (4) is equal to zero.

4. The strategy to test for unit roots follows Perron (1988).

5. Serletis (1992, 1994), used the procedure of Zivot and Adrews (1992).

6. We selected a VAR model with a constant to enter unrestricted following the procedure developed by Johansen (1992).

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