Inflation and Budget Deficit: What is the Relationship in Portugal?

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Abstract

The main causes of Portuguese inflation, based on annual data from 1954 to 1995, using the Johansen method, allows us to conclude that variation in Portuguese inflation is determined essentially by foreign inflation and by variation in the effective exchange rate of the Portuguese Escudo (PTE). In the long-term, the relationship between inflation rate and the growth rate of unit labour costs is almost unitary. However, the response of inflation change to the equilibrium error between inflation rate and changes in unit labour costs is slow and almost insignificant, while the response of unit labour costs to this disequilibrium is fast and significant, what suggests that the direction of causality is much more evident from the inflation rate on unit labour costs, than the reverse. The budget deficit as a percentage of GDP, are not significant in the short-term, in relation to variation in inflation as a dependent variable. However, it is significant in the relation to unit labour costs as a dependent variable, so we can have an indirect positive relation between inflation and lagged budget deficit.

Keywords: Inflation, Budget Deficit, Unit Roots, and Cointegration

JEL Classification: C12, C13, C32, E24, E31

Resumo

Os principais determinantes da variação da inflação no período 1954-95 parecem ser a inflação externa (ou a sua variação) e a variação da taxa de câmbio efectiva do escudo. Verifica-se uma relação de longo prazo entre a taxa de inflação e a taxa de variação dos custos unitários de trabalho quase unitária, mas a resposta da variação da inflação ao erro de equilíbrio entre a taxa de inflação e a variação dos custos unitários é lenta e quase insignificante ao passo que a resposta dos custos unitários de trabalho a esse desequilíbrio é rápida e significativa o que sugere que a direcção de causalidade é muito mais pronunciada da taxa de inflação para os custos de trabalho, do que ao contrário. Isto parece significar que os salários se ajustam imediatamente ao crescimento da inflação, enquanto a inflação se ajusta lentamente ao crescimento dos salários. O saldo orçamental em percentagem do PIB não é significativo na relação de curto prazo, na equação da inflação, no entanto, é significativo na equação dos custos unitários de trabalho, o que pode implicar relação positiva indirecta entre a variação da inflação e o défice orçamental desfasado.

1. Introduction

The relationship between the budget deficit and the inflation rate is not a stylized fact. In the economic literature there are at least two approaches, which try theoretically to establish a relation from budget deficit to inflation, but more recently some authors have empirically arrived to the relation from inflation to budget deficit.

In one approach, proposed by Sargent and Wallace (1981), it is assumed that the fiscal authority takes the measures without taking into account the current or future monetary policies. Thus, the monetary authority has to take restrictive measures in the short-term or in the long-term to defeat inflation. A restrictive monetary policy implies an increase in interest rate and the consequent reduction in product, giving rise to an increase in deficit *ceteris paribus* the fiscal policy. The fiscal authority will have to finance this increase in deficit, either by money emission, or by indebtedness. In the first case it implies an increase in inflation.

In another approach, inflation reduces the real stock of the public debt, thus public would tolerate an increase in inflation when the deficit is high because the public is adverse to an increase of the fiscal burden. However an increase in inflation, essentially the non-anticipated inflation, represents an inflationary tax.

Moreover, budget deficits also represent an additional aggregate demand that will give rise to an increase in inflation.

The economic literature has presented little empirical evidence of inflationary budget deficits. Santos (1992) analyses six countries of the European Union, where only three (including Portugal) seem to present inflationary deficits. Vieira (2000), also analyses six countries of the European Union (excluding Portugal), where it seems to exist more causal evidence from inflation to budget deficit, then in reverse.

The aim of this work is to analyse if the budget deficit constitutes one of the causes of inflation, inserted in a model that are looking for the main causes of the Portuguese inflation in the second half of the 20^{th} century, using annual data for the period 1954-1995. Thus, in section two an explicative model of the inflation will be considered, in section three we will present the chosen data and the reasons for their choice, in section four we will analyse the degree of stationarity of the used time series, in section five we will estimate the explicative model of the inflation considered in section two, using the method of Johansen to detect cointegration relations among the non stationary time series and applying the methodology of Rahbek and Mosconi (1999), which allows us to introduce

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stationary regressors in the VAR of cointegration through cumulated explanatory variables and simultaneously to use the trace or maximum eigenvalue tests. Finally in the sixth section we will present the main conclusions.

2. Model

The construction of a model is always a simplification of reality, given the multiplicity of variables that influence inflation, among them, an increase in the remuneration of productive factors, an increase in prices of imported products, a variation in the stock of money in circulation, a variation in the exchange rate, the budget deficit, expectations of inflation and the level and/or the variation in unemployment.

Considering the theory of mark-up, the monetarist theory of inflation and the possibility of the budget deficit being able to contribute to an increase in inflation, [1] we can consider the model:

$$\dot{P} = f\left(\dot{W} \overset{(+)}{-} \dot{Q},\ \overset{(+)}{\dot{P}}_M,\ \overset{(+)}{DEF},\ \dot{M} \overset{(+)}{-} \dot{y}\right) \tag{1}$$

$$\dot{P}_M \equiv \dot{P}_F + \dot{E} \tag{2}$$

Equation (1) contains the theory of mark-up where the firms set the price of their products above the marginal production cost. However, when the average cost is constant, it has been proved that the marginal cost is equal to the average cost, so that the prices (P) will be given by one mark-up above the average costs (CM):

$$P = \theta CM, \qquad \theta > 1 \tag{3}$$

If mark-up (θ) will be constant, the inflation rate ($\dot{P}$) will be equal to the rate of variation of the average costs. The average costs will vary in accordance with the wage variation corrected by the variation of the productivity ($\dot{W} - \dot{Q}$), which corresponds to the variation of the unit labour costs, and in accordance with the inflation imported in internal currency ($\dot{P}_M$). We assumed that the "other internal average costs" are constant.

Beyond the inflation for the costs, we also include in (1) the budget deficit in percentage of GDP (DEF) and the money growth beyond that necessary for transactions ($\dot{M} - \dot{y}$). In the inclusion of the budget deficit, one admits that an increase in government spending gives rise to inflation by demand, in virtue of the propensity of the government to spend being higher than the propensity of households to consume. The growth of money supply beyond that necessary for transactions, considering the income velocity of money to be constant, will have to imply an increase in inflation in accordance with the monetarist school.

[1] See Santos (1992) and Vieira (2000) on the relation between budget deficit and inflation.

The signals between parentheses on variables in equations (1) and (2) correspond to the signals expected for the coefficients of the relation. The equation (2) is an identity. The foreign inflation ($\dot{P}_F$) plus the variation of the effective indirect exchange rate ($\dot{E}$)[2] give the import inflation rate in terms of national currency. The aim of this work is to estimate the equation (1), where we will substitute the variable $\dot{P}_M$ for $\dot{P}_F$ and $\dot{E}$ in accordance with the equation (2) and we will try to see if budget deficit is inflationary.

3. Data

We use annual data whose justification in theoretical terms is given by Campbell and Perron (1991, p. 153) where, either due stationary analysis needs a long-term period, or because "seasonal adjustment procedures often create a bias toward nonrejection of a unit root hypothesis" (Campbell and Perron 1991, p. 153). In practical terms, it is difficult to get all the variables in quarterly terms in a compatible form for the study desired in period under consideration. However, this option is not exempt from problems either, because the majority of the available compatible series finish in 1995 and after 1995 they do not present a long enough number of observations as would be desirable for an econometrical study, so we opted to study the period 1954-95. As stated previously, we formulated the model on the basis of rates of change, so we opted to transform the available annual data into rates of change (with exception of the variable GG, which is a structure rate). Some authors think that the model would be richer if we used the original data, but we opted for rates of change because the variable that we intend to explain (the inflation rate) is generally $I(1)$, so it implies that the consumer price index (CPI) will be $I(2)$, and the model with variables $I(2)$ is not the aim of our study. Thus we selected seven annual variables for the period 1954-95, which we shall enumerate, presenting between square brackets its equivalence approached with variables of the theoretical model considered previously: P, inflation rate [$\dot{P}$]; ULC, rate of variation of the unit labour costs in firms [$\dot{W} - \dot{Q}$]; PM, rate of variation in import prices [$\dot{P}_M$]; E, nominal effective indirect exchange rate of the Escudo [$\dot{E}$]; PF, rate of variation in import prices in external currency [$\dot{P}_F$]; GG, General Government Balance in percentage of $GDPmp(cp)[-DEF]$; MY, rate of variation of the nominal stock of money ($M2^-$) corrected by the growth rate of the real $GDPmp[\dot{M} - \dot{y}]$.

These variables have been calculated from the *Historical Series for the Portuguese Economy* (1999) elaborated by Banco de Portugal, with the exception of the inflation rate (whose source is the annual CPI for the mainland, excluding housing rents, elaborated by Instituto Nacional de Estatística) and of the exchange rate (whose source is the statistical data of Mateus 1998).

[2] Indirect exchange rate means in terms of national currency, that $\dot{E} > 0 \Longleftrightarrow$ depreciation.

Once variables are selected, we will go on to study its stationarity; therefore the econometrical methodology to adopt in the estimation of the model formulated in the equation (1) depends on the stationarity degree of the time series.

The visual inspection (Figures 1-6 in Annex) points with respect to the stationarity of the foreign inflation (PF) with three outliers (1974, 1980 and 1986) which correspond to the effect of the first and the second oil-price shocks lagged by one year, as also applied to the favourable oil-price shock of 1985. The General Government Balance in percentage of the GDP seems to have suffered a structural break around the time of the revolution of April (between 1972 and 1974). Relative to the other variables, the visual inspection is not conclusive in terms of stationarity, although the inflation rate seems $I(1)$ as we expect from studies that some authors have carried out (see for example Cruz and Lopes 1999, p. 248). The exchange rate is practically constant up to 1974 due to the regimen of a fixed exchange rate, and has two very high peaks (1977, 1983) justified by high depreciation of the Escudo in periods of a high deficit in the Current Account,[3] with the aim of improving external competition.

4. Analysis of Stationarity of the Data

Visual inspection suggests the inexistence of a linear trend in the selected variables. However, we shall proceed, as if we did not know this from the starting point, in order to test the null hypothesis of a unit root. In the test on the existence of two unit roots, as we use the first differences of variables, it is enough to make the test on the model with a constant, because the visual inspection of the first differences of the selected variables indicates clearly the inexistence of any linear trend.

Thus, firstly we carried out tests on the existence of two unit roots (Table 1), secondly we carried out tests on the existence of a unit root (Table 2), and thirdly we carried out tests on the existence of a unit root in the time series under structural change with endogenous choice of the break point (Tb) (Table 3).

4.1. *Tests on the existence of two unit roots*

The Dickey and Pantula (1987) test allows us to reject the null hypothesis $H0 : I(2)$ against $I(1)$ in all variables studied to the level of significance of 1%, as we can see in Table 1. This test is based on the model:

$$\triangle^2 X_t = \mu + (\rho_1 - 1)X_{t-1} + (\rho_2 - 1)\triangle X_{t-1} + \sum_{i=1}^{k} \gamma_i \triangle^2 X_{t-i} + \varepsilon_t \qquad (4)$$

[3] Note that these two years precede agreements with the IMF to finance the Current Account deficit that had also reached two peaks.

where we test the null hypothesis of $I(2)$ against the alternative hypothesis of $I(1)$, that is: $H0 : \rho_1 - 1 = \rho_2 - 1 = 0$ against $Ha : \rho_1 - 1 = 0, (\rho_2 - 1) < 0$. To carry through this test we use the ratio t of $(\hat{\rho}_2 - 1)$ in the regression:

$$\triangle^2 X_t = \mu + (\rho_2 - 1)\triangle X_{t-1} + \sum_{i=1}^{k} \gamma_i \triangle^2 X_{t-i} + \varepsilon_t \tag{5}$$

and we use the critical values of the Dickey-Fuller table (Fuller 1976).

Table 1
Test on the existence of two unit roots

Variables	Dickey and Pantula (1987) test			
	k	$\tau_{\rho_2 - 1}$	LM(1) (F version)	Q(4)
P	3	-4.3814[a]	3.1880[.084]	1.515[.824]
ULC	0	-6.3341[a]	0.7275[.399]	6.510[.164]
E	2	-6.0322[a]	0.0028[.958]	0.069[.999]
PF	1	-6.6865[a]	1.9630[.170]	4.820[.306]
GG	1	-6.2116[a]	2.9966[.092]	3.511[.476]
MY	1	-7.6901[a]	0.7247[.400]	1.458[.834]

Notes: Model with a constant, annual data: 1954-1995.

[a] = significant at 1%; [b] = significant at 5%; [c] = significant at 10%.

The number of lags (k) of the second difference of each studied variable was selected, starting with $k - \max = 5$ and removing sequentially the last lag if insignificant at the 5% level until getting one lag that is significant in equation (4).

Tests LM and Q of Ljung-Box assure the absence of residual autocorrelation. Once the hypothesis of the existence of two unit roots is rejected at 1% level, we will test the hypothesis of the existence of one unit root.

4.2. *Tests on the existence of one unit root*

We applied the ADF test sequentially, starting with a model with a constant and a trend (CT) and selected k starting at $k - \max = 6$ and removed the last lag if insignificant at the 5% level until getting one lag that is significant (see Table 2). The three estimated models are of the form:

$$\text{Model 1 (CT)} : \triangle X_t = \mu + \beta(t - 1 - T/2) + (\rho - 1)X_{t-1} + \sum_{i=1}^{k} \gamma_i \triangle X_{t-i}$$

$$+\varepsilon_t \tag{6}$$

$$\text{Model 2 (C)} : \triangle X_t = \mu + (\rho - 1)X_{t-1} + \sum_{i=1}^{k} \gamma_i \triangle X_{t-i} + \varepsilon_t \tag{7}$$

$$\text{Model 3} : \triangle X_t = (\rho - 1)X_{t-1} + \sum_{i=1}^{k} \gamma_i \triangle X_{t-i} + \varepsilon_t \tag{8}$$

and the null hypothesis of existence of a unit root is $H0 : \rho - 1 = 0$.

We verified by the $LM(1)$ of Godfrey and the $Q(4)$ of Ljung-Box tests the absence of residual autocorrelation necessary to be able to apply the ADF test. We applied the joint tests Φ_3 and Φ_1 and the individual tests $\tau_{\beta\tau}, \tau_{\mu\tau}, \tau_{\mu\mu}$ of Dickey and Fuller (1981) to verify the existence of a trend or a constant in the case of the existence of a unit root, and thus we elaborated sequential tests until rejecting the null hypothesis of the existence of a unit root, in accordance with the advisable strategy for the use of the Dickey and Fuller tests described by Marques (1998, pp. 282–286). In the case of rejection of the existence of a unit root, we can test the existence of a trend or a constant using the traditional Student t test: in this case we present the p-value between square brackets in Table 2.

The joint and individual tests of Dickey and Fuller (1981), assuming from the outset that the unit root exists, are not very used in the practical way. It is more common to use visual inspection to see if a trend exists or not. In accordance with the individual test $\tau_{\beta\tau}$ (or $t_{\beta\tau}$ in the case of rejection of $H0$), we cannot reject the null trend as foreseen in the visual inspection, except for variable GG. Despite this result, we disagree that GG has a trend, in terms of visual inspection. Due to this discord, we initiated the selection of k in a model with a constant, and the variable GG is presented as $I(1)$ [Table 2: $GG(1)$ variable]. We think that this strange behaviour of GG is due to the structural break foreseen for visual inspection; therefore we will analyse it.

From the results of Table 2 we conclude that P and ULC are $I(1)$ and PF, E, GG and MY are $I(0)$.[4] Refering to Cruz and Lopes (1999), the fact of P to be $I(1)$ are in accordance with those authors.

4.3. *Tests for a unit root in time series under structural change with endogenous choice of the break point* (Tb)

Because of the hypothesis of structural break for variation of the mean in General Government balance in the percentage of $GDP(GG)$ we use the Perron and Vogelsang (1992) test. The break point (Tb) is endogenously selected by two processes: first, minimization of t statistic for testing $\alpha = 1[\min t_{\hat{\alpha}=1}]$, where α is the coefficient of the lagged variable to test the existence of a unit root; second, minimization of the $t_{\hat{\theta}}$ statistic (that is, t statistic for testing $\theta = 0$, where θ is the

4 Although the GG variable is presented as $I(1)$ [Table 2: line of $GG(1)$].

Table 2
ADF test on the existence of one unit root

Variables	ADF test								
	Mod.	K	$\tau_{\rho-1}$	Φ_3	Φ_1	$\tau_{\beta\tau}$	$\tau_{\mu\tau}$; $\tau_{\mu\mu}$	LM(1) F version	Q(4)
P	1 (CT)	4	0.1810	1.9547	-	-2.00035	0.75003	1.4327[.241]	1.0307[.905]
	2 (C)	4	-1.2200	-	0.7906	-	0.30231	3.3027[0.79]	1.2263[.874]
	3	4	-0.4035	-	-	-	-	3.5883[.068]	1.1935[.879]
ULC	1 (CT)	0	-2.9286	4.3813	-	-0.39471	0.09621	2.0579[.160]	3.1887[.527]
	2 (C)	0	-2.7556	-	3.8021	-	0.09724	0.7465[.393]	2.7305[.604]
	3	0	-1.6869	-	-	-	-	0.0488[.826]	5.4528[.244]
E	1 (CT)	1	-3.1178	5.0270	-	-0.51914	-0.06815	2.3556[.134]	1.8286[.767]
	2 (C)	1	-3.0771^{b}	-	4.7395c	-	1.5524	2.8316[.101]	2.369[.668]
	3	1	-2.6108^{b}	-	-	-	-	4.1643[.048]	5.2668[.261]
PF	1 (CT)	0	-4.0229^{b}	8.1044^{b}	-	-0.1610	1.3396	2.4627[.125]	3.0103[.556]
	2 (C)	0	-4.0740^{a}	-	8.3072^{a}	-	1.3566	2.5592[.118]	2.9995[.558]
	3	0	-3.8039^{a}	-	-	-	-	1.7010[.200]	3.2652[.514]
GG	1 (CT)	6	-4.0676^{b}	8.2772^{b}	-	-3.4250^{a}	-3.0674^{a}	0.0486[.945]	1.5873[.811]
	2 (C)	6	A	-	-	-	-	-	-
	3	6	A	-	-	-	-	-	-
GG(1)	2 (C)	7	-1.0715	-	0.7590	-	-0.60652	1.4605[.239]	1.3502[.853]
	3	7	-0.4856	-	-	-	-	1.5424[.226]	1.5606[.816]
MY	1 (CT)	0	-3.9708^{b}	8.0139^{b}	-	1.5896	3.4038^{a}	2.0737[.158]	3.7729[.438]
	2 (C)	0	-3.6045^{b}	-	6.4972^{b}	-	2.9550^{a}	4.1011[.050]	4.9019[.298]
	3	0	A	-	-	-	-		

Notes: Annual data: 1954-1995. Beginning of the tests in models with a trend, except $GG(1)$ where the tests begin in models with a constant. In $GG(1)$, because of there are residual autocorrelation in the former method, here we begin the selection with $k - \max = 10$, so $k = 7$.

A – we reject the null trend of a time series.

^{a} = significant at 1%; ^{b} = significant at 5%; ^{c} = significant at 10%.

coefficient of DU_t that represents the change in the mean of the time series) before one "crash" [$\min t_{\hat{\theta}}$] or maximization of the $t_{\hat{\theta}}$ statistic if we suspect an upward shift in the mean [$\max t_{\hat{\theta}}$].

In the first process, following the exposition of Perron (1997), we consider the choice of Tb in the all sample, although in the second process we restrict to the interval (0.15T, 0.85T), as suggested by Banerjee et al. (1992).

In the endogenous selection of k, we follow the first method described by Perron (1997, p. 359), which consists of a recursive procedure, where we started with $k - \max = 6$ and we eliminated lags successively not significant using two-sided t test at 10% level, to which Perron (1997) calls "t-sig" and which Perron and Vogelsang (1992, p. 313) consider leads to tests with higher power in almost all the studied cases.

In Table 3, we can observe the results of this test under the form of Innovational

Table 3
Tests for a unit root in time series under structural change

Series	Method	*IO* Model			*AO* Model		
		TB	k	$t_{\hat{\alpha}=1}$	TB	k	$t_{\hat{\alpha}=1}$
P	$\min \mathbf{t}_{\hat{\alpha}=1}$	1969	5	-1.61	1983	0	-1.99
	$\min \mathbf{t}_{\hat{\theta}}$	1983	4	-0.31	1989	5	-1.44
	$\max \mathbf{t}_{\hat{\theta}}$	1969	5	-1.61	1970	5	-0.92
ULC	$\min \mathbf{t}_{\hat{\alpha}=1}$	1971	1	-4.42[c]	1970	1	-4.46[c]
	$\min \mathbf{t}_{\hat{\theta}}$	1975	5	1.36	1989	0	-2.74
	$\max \mathbf{t}_{\hat{\theta}}$	1971	1	-4.42[b]	1972	5	-0.39
E	$\min \mathbf{t}_{\hat{\alpha}=1}$	1972	1	-4.05	1971	1	-4.10
	$\min \mathbf{t}_{\hat{\theta}}$	1985	3	-1.09	1988	3	-1.30
	$\max \mathbf{t}_{\hat{\theta}}$	1974	1	-4.03[c]	1975	6	-1.60
PF	$\min \mathbf{t}_{\hat{\alpha}=1}$	1973	0	-6.25[a]	1973	0	-6.29[a]
	$\min \mathbf{t}_{\hat{\theta}}$	1983	1	-4.61[b]	1984	0	-4.29[b]
	$\max \mathbf{t}_{\hat{\theta}}$	1970	1	-4.32[b]	1969	1	-4.40[b]
GG	$\min \mathbf{t}_{\hat{\alpha}=1}$	1972	6	-4.97[b]	1974	6	-5.34[a]
	$\min \mathbf{t}_{\hat{\theta}}$	1972	6	-4.97[b]	1974	6	-5.34[a]
	$\max \mathbf{t}_{\hat{\theta}}$	1961	6	-1.74	1987	3	-1.18
MY	$\min \mathbf{t}_{\hat{\alpha}=1}$	1967	0	-5.09[b]	1967	0	-5.21[a]
	$\min \mathbf{t}_{\hat{\theta}}$	1985	6	-0.20	1988	1	-2.42
	$\max \mathbf{t}_{\hat{\theta}}$	1967	0	-5.09[b]	1968	0	-4.57[a]

Notes: Sample period: 1954-95.

^{a,b,c} – Significant at 1, 5 and 10%, respectively.

$t_{\hat{\alpha}=1}$ in **bold**, means that we reject the existence of a unit root, at least at 5%.

Outlier (*IO*) and Additive Outlier (*AO*) Models. In *IO* model, the change of the series for the new structure becomes gradual, while in the *AO* model the change is sudden. The estimated equations are:

$$\text{IO Model}: \quad y_t = \mu + \theta DU_t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \triangle y_{t-i} + e_t \tag{9}$$

$$\text{AO Model}: \quad 1^{st} \text{ step}: y_t = \mu + \theta DU_t + \tilde{y}_t \tag{10}$$

$$2^{nd} \text{ step}: \tilde{y}_t = \sum_{i=0}^{k} w_i D(Tb)_{t-i} + \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} c_i \triangle \tilde{y}_{t-i} + e_t \tag{11}$$

The tests for structural change, either by the IO model or by the AO model, confirm the possibility of structural break for variable GG (from 1972 to 1974). This denotes an increase of the mean of GG gradually from 1972 to 1974 or instantaneously in 1974, this last year being most likely for the break.

Analysing the ADF and Perron and Vogelsang (1992) tests, we can say that the inflation rate (P) is $I(1)$ for all the tests and the rate of variation of the unit labour costs (ULC) is also $I(1)$ for almost all, so we must consider these two variables as $I(1)$ in the inflation model estimation, investigating the possibility of existence of relations of cointegration between them. The other variables, even with some doubts, are all considered $I(0)$, one of them (GG) with structural break (change in the mean) in accordance with the Perron and Vogelsang (1992) tests.

5. Estimation of an Explicative Model of the Inflation

We use the Johansen method as being the one that allows the detection of the presence of more than one cointegrating vector among variables in study.

There are stationary regressors in the VAR model, so we cannot use the critical values of Johansen (1996). Therefore, we follow the methodology of Rahbek and Mosconi (1999), which consists of adding to the VAR the cumulated explanatory $I(0)$ variables as $I(1)$ exogenous variables, and thus the critical values of the trace or eigenvalue tests of, among others authors, Pesaran et al. (2000) can be used. [5] First, as we have exogenous variables, the cointegrated VAR model to use corresponds to the conditional model: [6]

$$\triangle Y_t = \mu_c + \delta_c t + \sum_{i=1}^{k-1} \Psi_i \triangle X_{t-i} + \Pi_y X_{t-1} + \omega \triangle Z_t + \varepsilon_{ct} \tag{12}$$

where X_t is a $N \times 1$ vector of $I(1)$ variables, which we can partition into N_y endogenous $I(1)$ variables (Y_t) and N_z exogenous $I(1)$ variables (Z_t), such that $N_y + N_z = N$. Π_y is the long-run multiplier matrix of order $(N_y \times N)$ given by $\Pi_y = \alpha_y \beta'$, where α_y is a $(N_y \times r)$ matrix and β a $(N \times r)$ matrix of r cointegranting vectors.

[5] Referred to as PSS (2000), afterwards.

[6] We assume that the Z_t variables are weakly exogenous and they are not cointegrated between them, which implies that we can efficiently determine and test the parameters of long term (α and β), but with resource to the conditional model [see PSS (2000)].

The null hypothesis of the cointegration rank (existence of r cointegrating vectors) is written:

$$Hr : R[\Pi_y] = r, \qquad r = 0, \cdots, N_y \tag{13}$$

where "R" is the rank of the matrix.

First, in the estimation of the conditional model (12) we can consider 5 cases (or models) consonant with the restrictions imposed on the deterministic terms, following PSS (2000). Second, as we follow the methodology of Rahbek and Mosconi (1999), so $I(0)$ variables are included in $\triangle Z_t$ in equation (12) or in one of the 5 cases (models) consonant with the choice that is made. The cumulative sum of these $I(0)$ variables are $I(1)$ variables, corresponding to Z_t in the previous equation, enclosed therefore in X_t.

After this brief introduction we will try to estimate the corresponding model to the equation (1).

5.1. *Estimation of the long-term model*

In relation to the Model $P = f(ULC, PF, E, MY, GG)$, correspondent to equation (1) where we have two $I(1)$ variables (P and ULC) and four $I(0)$ variables (PF, E, MY and GG), we will apply the Methodology of Rahbek and Mosconi (1999) introducing the cumulated explanatory $I(0)$ variables into the cointegration relation and later we will test its exclusion from this relation using the likelihood ratio test. Thus, we will represent the model to study by:

$P\ ULC;\ csumPF\ csumE\ csumMY\ csumGG\&PF\ E\ MY\ GG$

where there are two endogenous $I(1)$ variables (P, ULC) and four exogenous $I(1)$ variables ($csumPF$, $csumE$, $csumMY$, $csumGG$) corresponding to the four $I(0)$ variables (PF, E, MY, GG), which are introduced into the short-term model. As we use the variable GG and not the variable DEF as in equation (1), the signal expected in the relation between P and GG will be negative, that is, when the budget deficit increases, the budget balance diminishes and one expects that the inflation rate will increase too.

In terms of k order of the VAR, we selected VAR(2), using either multivaried statistics, or univaried statistics so that the estimated residuals have no serial correlation (LB and LM tests), no autoregressive conditional heteroscedasticity (ARCH test) and they do not deviate too much from normality (BJ test), as Johansen (1996, p. 20) recommends.

With $k = 2$, whatever the model of the Johansen method is in terms of the deterministic terms, we cannot reject the existence of one cointegranting vector by the trace test, so we are going to choose the best model VAR(2) of cointegration in accordance with the deterministic terms considering $r = 1$. Following the PSS (2000) methodology, we cannot reject statistically the model IV.

Given VAR(2) and Model IV, one can confirm that it cannot reject the existence of one cointegranting vector, either by the trace test, or by the maximum eigenvalue test, as we can see in Table 4:

Table 4
Cointegration tests

Eigenvalue	Trace test			Maximal eigenvalue test		
λ	H0	Ha	Trace	H0	Ha	$\lambda\max$
0.61593	$r = 0$	$r \geq 1$	56.6284*	$r = 0$	$r = 1$	38.2773*
0.36794	$r \leq 1$	$r = 2$	18.3511	$r \leq 1$	$r = 2$	18.3511

* = significant at 5%.

The Schwarz Bayesian Criterion (SBC) also selects the model with $r = 1$. The vector normalized in relation to P (and identified) without restrictions with $X't = [P\ ULC\ csumPF\ csumE\ csumMY\ csumGG\ t]$ is given by:

$$\beta' = \begin{bmatrix} 1 & -1.2648 & 0.37018 & -0.002873 & -0.23262 & -0.16986 & 1.5306 \\ & (0.66047) & (0.35564) & (0.088995) & (0.16803) & (0.16352) & (1.1355) \end{bmatrix}$$

where one verifies that the cumulated variables have a relatively high standard error (between round brackets), and then it is probable that they are not significant in the long-term relationship. We cannot reject the hypothesis $H01 : \beta_3 = \beta_4 = \beta_5 = \beta_6 = 0$, by the likelihood ratio test with $\chi^2(4) = 4.0361[.401]$. And we cannot reject the joint test of $H01$ and trend=0 (hypothesis $H02 : \beta_3 = \beta_4 = \beta_5 = \beta_6 = \beta_7 = 0$), whose likelihood ratio test follows $\chi^2(5) = 4.5391[.475]$. So the long-term relationship is $P = 0.84016$ ULC and we have the cointegrating vector:

$$\beta' = \begin{bmatrix} 1 & -0.84016 & 0 & 0 & 0 & 0 & 0 \\ & (0.16427) & & & & & \end{bmatrix}$$

5.2. *Estimation of the short-term model*

The estimation of the multivaried model only with variables introduced initially in VAR(2) allows us to get the results in Table 5.

Analysing these equations, we verify that the variation of the inflation relates positively and significantly at 1% level to the foreign inflation and the variation of the exchange rate as expected, and relates to E_{t-1} and to MY_{t-1} at 10% level. The negative relation between $\triangle P$ and E_{t-1} means that $\triangle P$ relates positively to $\triangle E$ and the negative relation with MY_{t-1} (by the way, almost insignificant) is difficult to explain, but we will not worry about this, because in the parsimonious model, the Wald test suggests its exclusion from the model. The negative relation

Table 5
Estimation of the multivaried model

Equation	$\triangle P$	$\triangle$ ULC
No.observations/ regressors	T=40 [56-95]	T+40 [56-95]
Intercept	0.94720[0.276]	0.01434[0.991]
$\triangle P(-1)$	-0.21085[0.167]	0.09807[0.645]
$\triangle$ULC(-1)	-0.13663[0.154]	1.43280[0.163]
PF(-1)	-0.07557[0.341]	0.05518[0.662]
E(-1)	-0.25939[0.074]	-0.64378[0.003]
MY(-1)	-0.12325[0.097]	-0.19223[0.095]
GG(-1)	-0.21235[0.457]	-1.16320[0.007]
ECM1(-1)	-0.21484[0.136]	1.16730[0.000]
PF	0.39909[0.000]	0.31317[0.000]
E	0.43123[0.000]	-0.14700[0.315]
MY	-0.00749[0.918]	0.01192[0.908]
GG	0.38638[0.146]	0.65511[0.084]
$\bar{R}^2$	0.69	0.76
SEE	23.828	33.772
LM(1, 27)	0.28330[.599]	0.20973[.651]
RESET(1, 27)	0.03454[.854]	0.00066[.980]
BJ(2)	0.16744[.920]	0.03633[.982]
HET(1, 38)	0.73680[.396]	0.27639[.602]
ARCH(2, 26)	0.89157[.422]	0.81795[.452]

Note: See Annex about diagnostic tests description.

between $\triangle P$ and $\triangle ULC_{t-1}$ can be explained by Santos (1992) conclusion about the positive relationship between inflation and variation in unit labour costs lagged by two periods. We can say that if the coefficient of the positive relation between $\triangle P$ and ULC_{t-2} is higher than the coefficient of the positive relation between $\triangle P$ and ULC_{t-1}, we will have a negative coefficient in the relation between $\triangle P$ and ULC_{t-1} because, $\triangle ULC_{t-1} = ULC_{t-1} - ULC_{t-2}$. So, the observed relationship in our model (however not significant) could mean that the inflation responds to lagged costs as Santos (1992) concludes.

The long-term relationship $(P - 0.84016ULC)$ represented by $ECM1$ presents the expected signal but in this initial model this is not significant. This strengthens the weak exogeneity of the inflation rate in this model. The relationship between $\triangle P$ and GG hasn't the expected signal and isn't significant, but the relationship between $\triangle P$ and GG_{t-1} has the expected signal.

The ULC variation relates positively and significantly at 1% to $ECM1_{t-1}$ with a coefficient close to one. This means that the labour costs responds fast and significantly to an increase in inflation above increase of the labour costs in the last period. We can explain this by trade union pressing to increase wages in the next period. $\triangle ULC$ relates also positively and significantly at 1% to PF and negatively at 1% to E_{t-1}. PF is the proxy to inflationary expectations and E_{t-1} is the proxy to $\triangle U_{t-1}$, as the author has studied in an earlier work.

$\triangle ULC_t$ relates negatively and significantly at 1% to GG_{t-1}. How can we explain this? As the negative relation with GG_{t-1} means positive relation with budget deficit in percentage of GDP, we can say that, *the high budget in the last period could increase the inflationary expectations*, which would imply some pressing in wage increase. So, the general government balance (GG) can indirectly exert influence upon inflation through $\triangle ULC$.[7] However, for this we need a significant $ECM1$ at $\triangle P$ equation. So, the budget deficit can have a positive influence upon inflation through costs, instead of through demand as we had supposed in the model *ab initio* (Section 2).

The diagnostic tests indicate that the residuals are not autocorrelated, are homoscedastics, normal and we cannot reject correct specification of the model. The autoregressive conditional heteroscedasticity is also absent until the second order.

In $\triangle P$ equation, all the residuals are inside the line bands of double standard deviation and CUSUM and CUSUMSQ tests do not cross any of the significant bars at 5% level.

We tried to remove from the equation of $\triangle P$ the variables that were not significant at the 10% level, using the Wald test on the joint nullity of its coefficients, to reestimate **parsimonious equations**. The Wald test does not allow us to reject all the non-significant variables, so, after some attempts we kept $PF_{t-1}, ECM1_{t-1}$ and ULC_{t-1} in the regression of $\triangle P$, despite its non-significance in the initial regression. As E and E_{t-1} have symmetrical coefficients, we substitute them for $\triangle E$.

At the parsimonious equation DP1 (Table 6), foreign inflation (PF), lagged foreign inflation (PF_{t-1}) and the variation of exchange rate ($\triangle E$) are significant at 1%, $ECM1_{t-1}$ becomes significant at 2% and $\triangle ULC_{t-1}$ is significant at 8%. Reestimating the previous equation for 1956-88 (equation DP2), we cannot reject, either the predictive capacity after-1988 or the structural stability before and after 1988, using the Chow (1960) tests.

The introduction of dummies[8] (equations DP3 to DP5) allows us to verify

[7] Increase in GG would imply diminishing in $\triangle ULC$, and this would diminish inflation.

[8] $Dum74$ (value 1 in 1974 – first oil shock and April Revolution), $Dum79$ (value 1 in 1979 – second oil shock), $Dum80$ (value 1 in 1980 – Escudo Revaluation), $Dum87$ (value 1 in 1987 – favourable external conjuncture), NS (value 1 up to 1973 – New State), EEC (value 1 after 1986 – Member of the EEC),

that there are three of them significant (Dum87, Dum80 and EMS), which contribute to diminish inflation. Besides, with the introduction of dummies, ECM_{t-1} becomes significant at 1%. The parsimonious model with three dummies together (equation DP5) implies that almost all variables are significant at 1%. The dummy Dum80 has a strong impact on diminishing the coefficient of $\triangle E$ and the dummy EMS has also some impact on that, what we were expecting, because the former refers to revaluation of Escudo and the latter refers to the period after 1992 (participation in the Exchange Rate Mechanism of the European Monetary System), when Escudo devaluated. The dummy EMS also changes the coefficient of $ECM1_{t-1}$ from about 0.23 to 0.27, but the parsimonious estimative without this dummy gives a coefficient of $ECM1_{t-1}$ close to that estimated in original model.

Table 6
Parsimonious Equations of $\triangle P$

Equation/ regressors	DP1 $T=40$ [56-95]	DP2 $T_1=33, T_2=7$ [56-88]	DP3 $T=40$ [56-95]	DP4 $T=40$ [56-95]	DP5 $T=40$ [56-95]
Inpt	.19491[.675]	.41079[.438]	.50905[.260]	.46775[.275]	.95613[.033]
△ULC(-1)	-.15886[.076]	-.16607[.085]	-.14446[.082]	-.17189[.033]	-.19766[.010]
ECM1(-1)	-.22219[.014]	-.23064[.015]	-.23467[.006]	-.22794[.005]	-.26882[.001]
PF	.36942[.000]	.36140[.000]	.38038[.000]	.40473[.000]	.39723[.000]
PF(-1)	-.17747[.006]	-.19457[.003]	-.23049[.001]	-.20105[.002]	-.21409[.000]
△E	.34630[.001]	.38018[.001]	.35160[.000]	.27872[.003]	.25365[.004]
Dum80	-	-	-	-5.7270[.037]	-5.5815[.027]
Dum87	-	-	-6.3715[.014]	-5.9205[.017]	-6.7448[.004]
EMS	-	-	-	-	-2.8900[.012]
$\bar{R}^2$	.67763	.73302	.72366	.75193	.79159
SEE	2.4218	2.3823	2.2422	2.1244	1.9472
DW	2.3074	2.6018	2.2457	2.0534	2.4599
LM(1, T-k-1)	1.0944[.303]	3.4361[.075]	.69820[.410]	.062580[.804]	2.6444[.114]
$\text{RESET}_{(1,T-k-1)}$	.22702[.637]	.47222[.498]	.28806[.595]	.43092[.516]	.45778[.504]
BJ(2)	1.7416[.419]	1.3455[.510]	.99786[.607]	.30506[.859]	.066843[.967]
HET(1, T-2)	.53309[.470]	.45548[.505]	.85361[.361]	.73029[.398]	.92964[.341]
$\text{ARCH}_{(2,T-k-2)}$	.85107[.436]	.51264[.605]	.78752[.464]	1.1698[.324]	1.4784[.245]
$\text{Chow}_{(T2,T1-k)}$	-	1.1625[.356]	-	-	-
$\text{Cov}_{(k,T1+T2-2k)}$	-	1.2904[.294]	-	-	-

EMS (value 1 after 1992 – Participation in the ERM of the EMS).

Table 6: Parsimonious Equations of $\triangle P$ (continuation)

Equation/ regressors	DP6 $T = 22$ [74-95]	DP7 $T = 22$ [74-95]	DP8 $T_1 = 15, T_2 = 7$ [74-88]	DP9 $T = 22$ [74-95]	DP10 $T = 22$ [74-95]	DP11 $T = 22$ [74-95]
Inpt	-.4107[.605]	-.84042[.220]	-.44478[.715]	.24951[.747]	.21694[.763]	1.4552[.078]
△ULC(-1)	-.09785[.449]	-	-.07533[.639]	-.06354[.586]	-.10580[.346]	-.13940[.155]
ECM1(-1)	-.14045[.233]	-	-.11875[.420]	-.16019[.137]	-.15937[.115]	-.24099[.015]
PF	.39456[.000]	.39254[.000]	.38827[.000]	.40598[.000]	.43159[.000]	.41830[.000]
PF(-1)	-.21178[.020]	-.24943[.001]	-.23474[.030]	-.28474[.003]	-.24882[.006]	-.28112[.001]
△E	.41869[.003]	.51320[.000]	.48949[.011]	.42585[.001]	.34706[.006]	.29909[.006]
Dum80	-	-	-	-	-5.5659[.086]	-5.4774[.050]
Dum87	-	-	-	-6.6181[.043]	-6.1142[.046]	-8.0831[.006]
EMS	-	-	-	-	-	-3.3768[.022]
$\bar{R}^2$	.73028	.73721	.78242	.78294	.81304	.86739
SEE	2.8666	2.8295	3.0902	2.5716	2.3866	2.0100
DW	2.3225	2.4867	2.8830	1.9762	1.6512	2.3613
LM(1, T-k-1)	.55716[.467]	1.1966[.289]	2.9110[.126]	.001950[.965]	.55247[.471]	.80555[.387]
$\text{RESET}_{(1,T-k-1)}$	.17906[.678]	.68751[.419]	.23013[.644]	.52779[.480]	.95560[.346]	2.1592[.167]
BJ(2)	1.1444[.564]	1.3126[.519]	.34645[.841]	.95970[.619]	.18871[.910]	.88520[.642]
HET(1, T-2)	1.5600[.226]	2.0063[.172]	2.2415[.158]	1.2257[.281]	.53778[.472]	1.3512[.259]
$\text{ARCH}_{(2,T-k-2)}$	2.1719[.151]	4.6314[.026]	.58628[.582]	.19414[.826]	.22891[.799]	.020099[.980]
$\text{Chow}_{(T2,T1-k)}$	-	-	.68115[.687]	-	-	-
$\text{Cov}_{(k,T1+T2-2k)}$	-	-	.79736[.593]	-	-	-

Notes: Dependent Variable: $\triangle P$; Estimation Method: OLS; ECM1= P – 0.84016ULC estimated on model: $P\ ULC$; $csumPF$, $csumE$, $csumMY$, $csumGG\&PF\ E\ MY\ GG$. See Annex about diagnostic tests description. Between square brackets: p-value or sample period (on the top). On the estimated coefficients, the null hypothesis is $H0 : \beta = 0$, and the Student t test is used.

The coefficients of $\triangle ULC_{t-1}$ and of $ECM1_{t-1}$ becomes not significant and with lower absolute value in the period 1974-95 without dummies (equation DP6). Otherwise, the coefficient of $\triangle E$ increases. However, the exclusion of the variables that were not significant (equation DP7) generates autoregressive conditional heteroscedasticity, so that we opted to keep these two variables. In the period 1974-95, we cannot also reject either the predictive capacity after-1988 or structural stability before and after 1988 (equation DP8).

The dummies Dum80, Dum87 and EMS are also significant in period 1974-95 and $ECM1_{t-1}$ becomes significant at 5% (with the three dummies in equation DP11), but $\triangle ULC_{t-1}$ is always not significant in this period. The comparison of the period 1974-95 (equation DP11) with the period 1955-95 (equation DP5) allows us to notice the increase of absolute value of the coefficients of $PF, PF_{t-1}, \triangle E$ and Dum87, in opposite to the diminishing of the absolute value of the coefficients of ECM_{t-1} and ULC_{t-1}. This highlight the increase of importance that foreign inflation and variation of exchange rate have in determining the domestic inflation with the openness of Portuguese economy after April revolution, together with political measures to increase competition abroad (Escudo devaluation).

In the parsimonious model of $\triangle P$*, the equilibrium error* $(ECM1_{t-1})$ *is significant, so this support the possibility of the negative relationship between* $\triangle ULC_t$ *and* GG_{t-1}*, (equation* $\triangle ULC$*), which can indirectly be able to influence the inflation, as we said earlier. There would be positive relationship between inflation and budget deficit, however lagged by two periods: when GG diminishes (budget deficit increases) in period* $t-2$*, implies that* $\triangle ULC_{t-1}$ *increases (by equation* $\triangle ULC$*), and possibly we will have* $P_{t-1} < ULC_{t-1}$*, (that is,* $ECM1_{t-1} < 0$*), implying an increase in* $\triangle P_t$ *(by equation* $\triangle P$*).*

6. Conclusions

The main causes of the variation in inflation in the period 1954-95 seem to be foreign inflation (or its variation) and the variation in the effective exchange rate of the Escudo. There is a long-term relationship between the inflation rate and the growth rate of unit labour costs, almost unitary, but the response of the variation in inflation to the equilibrium error between the inflation rate and the variation in unit labour costs is slow and almost insignificant, while the response of unit labour costs to that disequilibrium is fast and significant, which suggests that the direction of causality is much more evident from the inflation rate to the unit labour costs, than the reverse. This seems to mean that the wages adjust to growth in inflation quickly, while inflation adjusts to growth in wages slowly.

The variation in nominal money stock, corrected by the growth rate of the real GDP, as well as the general government balance in percentage of GDP, are not significant in the short-term relationship, in inflation equation, however, the general government balance is significant in unit labour costs equation, so this can imply a positive and indirect relationship between inflation and lagged budget deficit.

The comparison of our results with those of other authors allows us to verify that our conclusions are not very different to those of the majority of the authors who have made studies for the 1970s and 1980s, so that one sub-period strongly influences our conclusions. Santos (1992) concludes that the budget deficit seems to be inflationary, but only in 50% of the analysed countries, among them Portugal, and Vieira (2000) concludes that there is little support for the idea that budget deficits have contributed to inflation in the majority of European countries,[9] so therefore we do not find strange our conclusion in relation to the non-influence of the budget deficit on the variation in inflation.

[9] There is more evidence so that in its model the inflation has contributed for deficits.

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Annex

Variables Plots

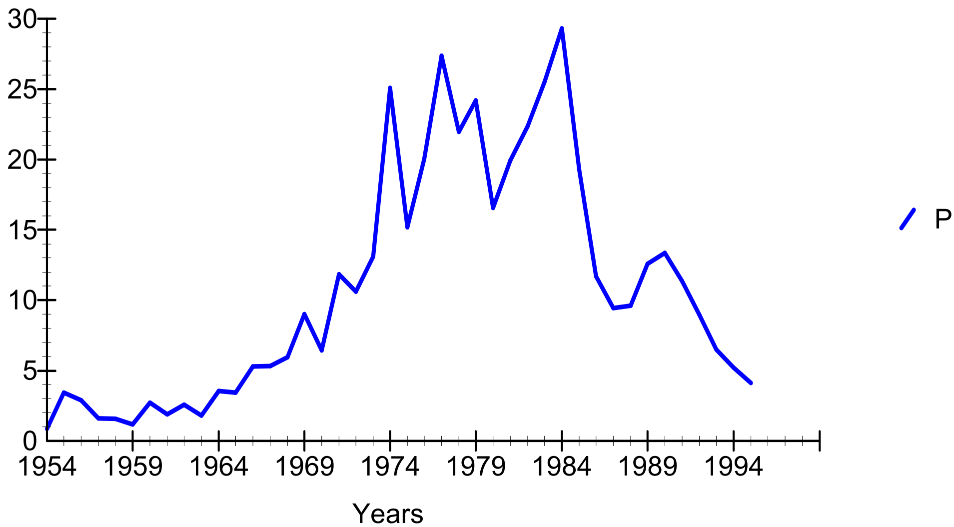

Fig. 1. Inflation rate

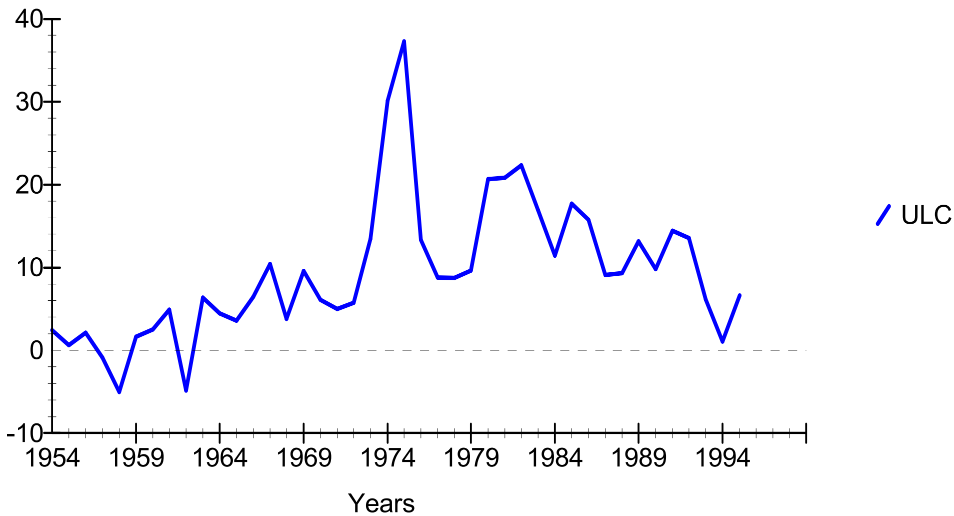

Fig. 2. Rate of variation of the unit labour costs in firms

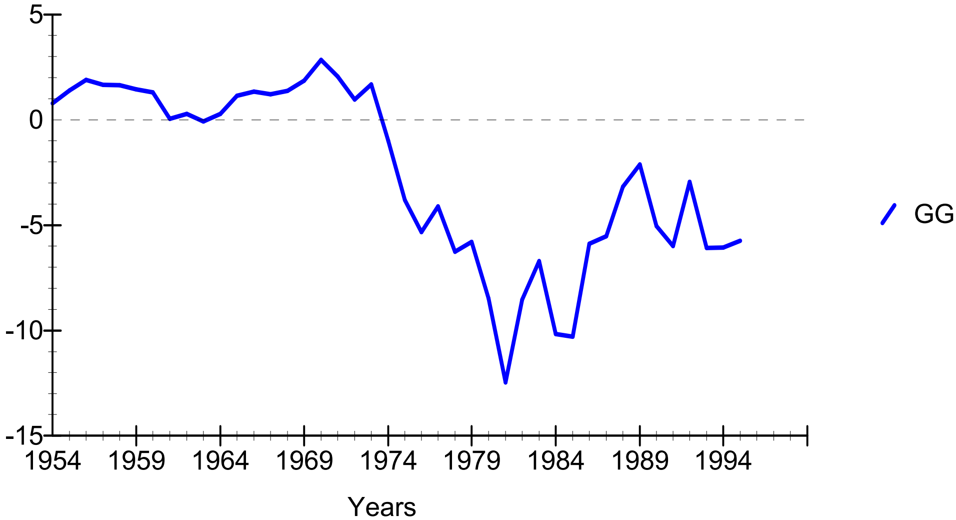

Fig. 3. General government balance in percentage of GDPmp(cp)

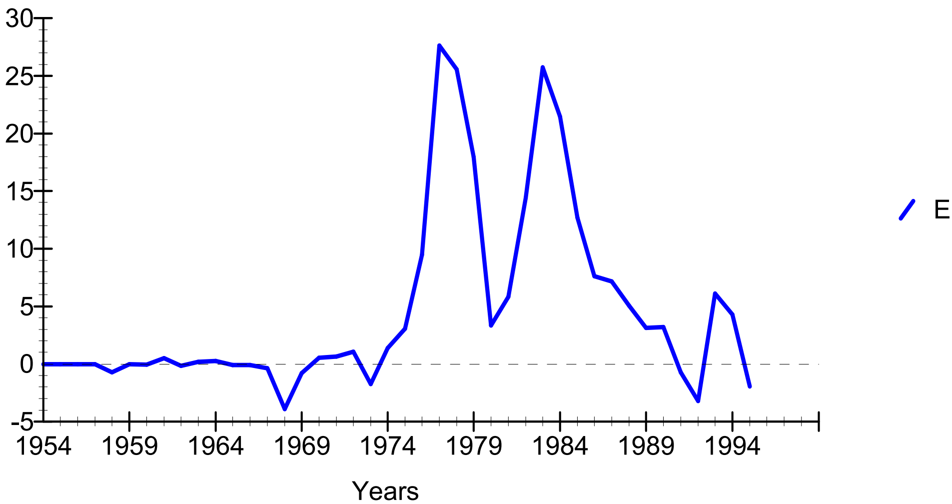

Fig. 4. Nominal effective indirect exchange rate of the Escudo

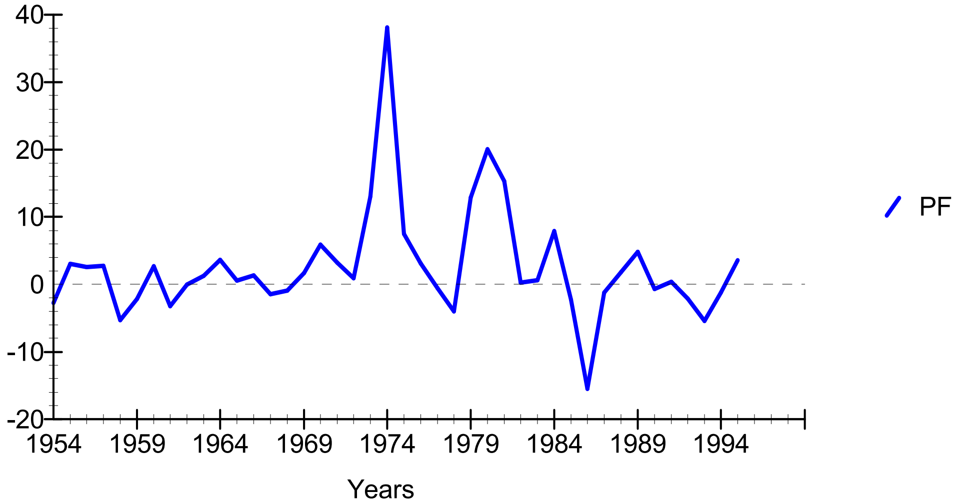

Fig. 5. Rate of variation in import prices in foreign currency

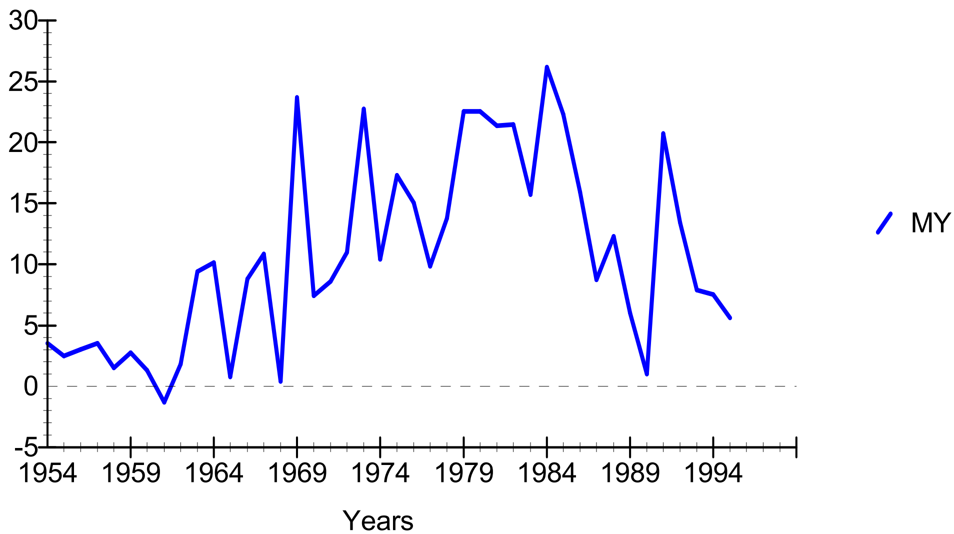

Fig. 6. Rate of variation of the nominal stock of money (M2) corrected by the growth rate of the real GDPmp

Notes about Diagnostic Tests

Diagnostic tests: We use the F version of diagnostic tests because Marques (1998) citing Kiviet (1986) [10] said that in small samples the F version is preferable. In BJ test we present the LM version following a $\chi^2(2)$, because the F version does not apply in this test. The degrees of freedom of the F test are in round brackets, which depend on the k and T: **T**=number of observations used in regression; **k**=number of estimated coefficients; $\mathbf{T_1}$=sub-sample used in estimation; $\mathbf{T_2}$=Period post-sample (forecasting test) or second sub-sample (stability test, only possible when $T_1 > k$ and $T_2 > k$).

Diagnostic tests description:

LM – statistic of Lagrange Multiplier test for serially correlated residuals [based in Godfrey(1978)]. [11]
RESET – statistic of Ramsey(1969)'s [12] RESET test of functional form misspecification.
BJ – statistic of Jarque-Bera's test of normality of regression residuals [based in Bera and Jarque(1981)]. [13]
HET – statistic of Heteroscedasticity test [see Pesaran and Pesaran (1997)].
ARCH – statistic of Autoregressive Conditional Heteroscedasticity test [Engle (1982)'s [14] test].
Chow – statistic of Predictive failure test (2^{nd} test of Chow (1960)).
Cov – statistic of Chow's test of stability of regression coefficients (1^{st} test of Chow (1960)).

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