Causes of Portuguese Inflation: An Econometric Application

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ABSTRACT

A study of the 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.

Because of the fact that the variation in nominal money stock, corrected by the growth rate of real GDP, as well as budget deficit as a percentage of GDP, are not significant variables in the short run, in relation to variation in inflation as a dependent variable, we can conclude that inflation is caused essentially by costs. The costs that are highly significant in the short run are those created by the rate of variation in import prices (determined either by foreign inflation or by variation in the effective exchange rate).

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Introduction

The causes of Portuguese inflation was the focus of the study of some authors in the last quarter of the 20th century, including Abel Mateus (1980), Jose Girão (1984), Robalo Marques (1990), Jorge Santos (1992) and more recently Robalo Marques (1995), Cunha and Machado (1996) and Catela Nunes (1998).

The aim of this work is to give continuity to these studies, looking for the main causes of Portuguese inflation in the second half of the 20th century, using annual data for the period 1954-1995. Thus, in the first stage an explicative model of the inflation will be considered, in the second stage we will present the chosen data and the reasons for their choice, in the third stage we will analyze the degree of stationarity of the used time series and in the fourth stage we will estimate the explicative model of the inflation considered in the first stage, using the method of Johansen to detect cointegration relations among the non-stationary series and applying the methodology of Rahbek and Mosconi (1999), which allows us to introduce stationary regressors in the VAR of cointegration through cumulated explanatory variables and simultaneously to use the trace or maximum eigenvalue tests. Finally in the fifth stage we will present the main conclusions.

1. The Considered 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 the 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 Phillips curve, the theory of mark-up, the monetarist theory of inflation[1] and the possibility of the budget deficit being able to contribute to an increase in inflation,[2] we can consider the model:

$$\dot{P} = f\left(\overset{(+)}{\dot{W} - \dot{Q}},\ \overset{(+)}{\dot{P}_M},\ \overset{(+)}{D\dot{E}F},\ \overset{(+)}{\dot{M} - \dot{y}}\right) \qquad [1.1]$$

[1] See Surrey (1989). This author presents the model of inflation for the costs (theory of mark-up) separate from the monetarist model of the inflation. We have joined the two theories in a single model.
[2] See Jorge Santos (1992) and Carlos Vieira (2000) on the relation between budget deficit and inflation.

$$\dot{W} - \dot{Q} = g\left(\overset{(-)}{U}, \overset{(+)}{\dot{P}^e}\right) \qquad [1.2]$$

$$\dot{P}_M \equiv \dot{P}_F + \dot{E} \qquad [1.3]$$

Equation (1.1) 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):

$$P = \theta\, CM, \qquad \theta > 1 \qquad [1.4]$$

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 imported inflation in domestic currency ($\dot{P}_M$).[3]

Beyond the inflation for the costs, we also include in (1.1) the inflation for monetary emission beyond that necessary for transactions ($\dot{M} - \dot{y}$) and the budget deficit in percentage of GDP (DEF). 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. In the inclusion of the budget deficit, one admits that an increase in public consumption gives rise to inflation by demand, in virtue of the propensity of the government to consume being greater than the propensity of households to consume.

The inclusion in the model of the two variables (budget deficit and variation in the nominal stock of money) raises some problems, since there are periods where the government uses the monetary emission to finance its deficit, so there is the possibility of correlation between these variables. Thus, before analyzing this complete model, we studied two sub-models, one without deficit,[4] and another one without variation of money supply. As the basic conclusions do not change, we will study the complete model. The signals between parentheses on variables in equations (1.1) and (1.2) correspond to the signals expected for the coefficients of the relation.

[3] It is assumed that the "other internal average costs" are constant. See Agostinho Rosa (2000).
[4] See Agostinho S. Rosa (2003).

The equation (1.2) corresponds to the augmented Phillips curve with expectations, considering that growth in wages is positively related to growth in productivity ($\dot{Q}$) in accordance with Burda and Wyplosz (1993, p. 245).

The equation (1.3) is an identity. The foreign inflation ($\dot{P}_F$) plus the variation of the effective indirect exchange rate ($\dot{E}$)[5] give the import inflation rate in terms of domestic currency. The aim of this article is to estimate the equation (1.1), where we will substitute the variable $\dot{P}_M$ for $\dot{P}_F$ and $\dot{E}$ in accordance with the equation (1.3).

2. 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 the 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.[6] 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.[7] 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 I(2) variables is not the aim of our study. Thus we selected eight annual variables for the period 1954-95, which we shall enumerate, presenting between square brackets its approached equivalence with variables of the theoretical model considered previously: P, inflation rate [$\dot{P}$]; U, unemployment rate (board sense) [U]; CTUPEV, 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,

[5] Indirect exchange rate means in terms of domestic currency, that $\dot{E} > 0 \Leftrightarrow$ depreciation.
[6] And we use data of the *Historical Series for the Portuguese Economy*, Bank of Portugal, 1999, which finished in 1995.
[7] With the exception of the variable SPA, which is a structure rate.

rate of variation in import prices in foreign currency [$\dot{P}_F$]; SPA, 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 the Portuguese Central Bank, 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, INE) and of the exchange rate (whose source is the statistical data of Abel Mateus(1998)).

Once variables are selected, we will study their stationarity, and the econometrical methodology to adopt in the estimation of the model formulated in the equation (1.1) depends on the order of integration of the time series.

The plots of variables (visual inspection) points to the stationarity of 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, and to the favourable oil-price shock of 1985. The unemployment rate (U) seems to have suffered a structural break around 1974/75, which corresponds to the revolution of April. The General Government Balance in percentage of the GDP also 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 inflation rate seems I(1), as we expect from studies that some authors have carried out.[8] The exchange rate is practically constant up to 1974 due to the regimen of a fixed exchange rate,[9] 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,[10] with the aim of improving external competition.

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[8] See for example Cruz and Lopes (1999, p. 248).
[9] See Botas and Sousa (1995, p. 14).
[10] Note that these two years precede agreements with the IMF to finance the Current Account deficit that had also reached two peaks.

3. Analysis of stationarity of the data

Firstly we carried out tests on the existence of two unit roots (Table I - Annex), secondly we carried out tests on the existence of a unit root (Table II - Annex), 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 III - Annex).

3.1 Tests on the existence of two unit roots

The Dickey and Pantula (1987) test allows us to reject the null hypothesis I(2) against I(1) in all variables studied to the level of significance of 1%, as we can see in Table I of the Annex (1^{st} step). 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 this test, 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.

Tests LM and Q of Ljung-Box assure the absence of residual autocorrelation.

The second step of the test of Dickey and Pantula to test H0: I(1) against I(0) only rejects H0 for variables CTUPEV and PF. The rejection of H0 for the variable CTUPEV is stranger because, in the ADF test for the existence of a unit root, it is not rejected, as we will see.[11] Once the hypothesis of the existence of two unit roots is rejected, we will test the hypothesis of the existence of one unit root.

3.2 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 II - Annex).

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

[11] And the same result happens in others tests not presented here, as PP and KPSS.

Dickey and Fuller tests described by Robalo 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 II of the Annex.

The joint and individual tests of Dickey and Fuller (1981), assuming from the outset that a unit root exists, are not commonly used. It is more common to refer to the 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 SPA. Despite this result, we disagree that SPA 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 SPA is presented as I(1) [Table IIA-Annex]. We think that this strange behaviour of SPA is due to the structural break foreseen for visual inspection; therefore we will analyze it.

From the results of Table II we conclude that P, CTUPEV and U are I(1) and PF, E, MY and SPA are I(0).[12] Refering to Cruz and Lopes(1999), the fact of U and P being I(1) is in accordance with those authors. Cruz and Lopes(1999, p. 248) also raise doubts relative to the nominal stock of money being I(2), opting to considering it I(1), which is in accordance with our result of the rate of variation of the nominal stock of money corrected for the rate of growth of the product being I(0), since the growth rate of the real GDPmp is clearly a I(0) variable.[13]

3.3 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 (SPA) and in the unemployment rate (U), 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 α=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 coefficient of DU_t that represents the change in the mean of the time

[12] Although the SPA variable is presented as I(1) in the Table IIA.

[13] Although we do not present it here, the rate of variation of the nominal stock of money is also rejected as I(1), as well as the rate of growth of the GDPmp.

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 whole sample, although in the second process we restrict it 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 a two-sided t test at 10 % level, which Perron(1997) calls "t-sig" and which Perron and Vogelsang(1992, p. 313) consider leads to tests with greater power in almost all the studied cases.

In Table III (Annex), we can observe the results of this test under the form of Innovational Outlier (IO) and Additive Outlier (AO) Models. In the IO model, the change of the series for the new structure becomes gradual, while in the AO model the change is sudden. The tests for structural change, either by the IO model or by the AO model, confirm the possibility of a structural break for the unemployment rate (from 1973 to 1975), and for variable SPA (from 1972 to 1974).[14] This denotes an increase of the mean of U gradually from 1973 to 1975 or instantaneously in 1975, this last year being most likely for the break in accordance with Cruz and Lopes(99); the same with the mean of SPA from 1972 to 1974.

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 (CTUPEV) 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), the two of them (U and SPA) with structural break (change in the mean) in accordance with the Perron and Vogelsang (1992) tests.

However, as the rejection of I(1) in the SPA with breaking in 1974 <u>is significant at 1% by two methods of selection</u> of the point of breaking (Tb) and the rejection of I(1) in the unemployment rate with breaking in <u>1975 is only significant at 5% by one method of selection</u>, we can admit that U is I(1) and that SPA is I(0).

[14] Note that the first point of breaking corresponds to the IO model and the second to the AO model.

4. 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 other authors, Pesaran, Shin and Smith(1999) can be used.[15] First, as we have exogenous variables, the cointegrated VAR model to use corresponds to the conditional model:[16]

$$\Delta Y_t = \mu_c + \delta_c\, t + \sum_{i=1}^{k-1} \Psi_i \Delta X_{t-i} + \Pi_y X_{t-1} + \omega \Delta Z_t + \varepsilon_{ct} \qquad [4.1]$$

where X_t is a $N\times 1$ vector of I(1) variables, which we can divide 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.

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

$$\text{Hr: R } [\Pi_y] = r, \qquad r = 0, \ldots, N_y; \qquad [4.2]$$

where "R" is the rank of the matrix.

In the estimation of the conditional model (4.1) we can consider 5 cases (or models) consonant with the restrictions imposed on the deterministic terms. Following PSS(99) we have:[17]

Case I (No intercepts; no trends):

$$\mu_c = \delta_c = 0 \Rightarrow \Delta Y_t = \sum_{i=1}^{k-1} \Psi_i \Delta X_{t-i} + \Pi_y X_{t-1} + \omega \Delta Z_t + \varepsilon_{ct} \qquad [4.3]$$

[15] Referred to as PSS(99), afterwards.

[16] 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(99)].

[17] It corresponds to the 5 cases considered in the program Microfit 4.0. On the differences in cases III and V relative to models 3 and 5 of Johansen (1996), when it does not have exogenous variables, see PSS(99). It is also useful to see Mackinnon et al.(1999, p. 568) which compares the 5 cases of PSS(99) with tables

Case II (Restricted intercepts; no trends):

$$\begin{cases}\mu_c = -\Pi_y \eta \\ \delta_c = 0\end{cases} \Rightarrow \ \Delta Y_t = \sum_{i=1}^{k-1} \Psi_i \Delta X_{t-i} + \Pi_y^{*}\left(X_{t-1}^{'},1\right)^{'} + \omega \Delta Z_t + \varepsilon_{ct} \qquad [4.4]$$

where $\Pi_y^{*} = \Pi_y\left(I_N, -\eta\right)$ with I_N = identity matrix (N×N).

Case III (Unrestricted Intercepts; no trends):

$$\begin{cases}\mu_c \neq 0 \\ \delta_c = 0\end{cases} \Rightarrow \ \Delta Y_t = \mu_c + \sum_{i=1}^{k-1} \Psi_i \Delta X_{t-i} + \Pi_y X_{t-1} + \omega \Delta Z_t + \varepsilon_{ct} \qquad [4.5]$$

Case IV (Unrestricted intercepts; restricted trends):

$$\begin{cases}\mu_c \neq 0 \\ \delta_c = -\Pi_y \gamma\end{cases} \Rightarrow \ \Delta Y_t = \mu_c + \sum_{i=1}^{k-1} \Psi_i \Delta X_{t-i} + \Pi_y^{**}\left(X_{t-1}^{'},t\right)^{'} + \omega \Delta Z_t + \varepsilon_{ct} \qquad [4.6]$$

where $\Pi_y^{**} = \Pi_y\left(I_N, -\gamma\right)$.

Case V (Unrestricted intercepts; unrestricted trends):

$$\begin{cases}\mu_c \neq 0 \\ \delta_c \neq 0\end{cases} \qquad \Rightarrow \text{The model of the equation (4.1) will be estimated.}$$

First, these 5 cases are elaborated for $N_z>0$ (existence of weakly exogenous variables), but give results for $N_y=N$ as a special case when $N_z=0$ (inexistence of weakly exogenous variables). Second, as we follow the methodology of Rahbeck and Mosconi(1999), our I(0) variables are included in ΔZ_t in equation 4.1 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[18] we will try to estimate the corresponding model to the equation (1.1).

4.1 Estimation of the Long –Term Model

In relation to the Model P=f(CTUPEV, PF, E, My, SPA), correspondent to equation 1.1 where we have two I(1) variables (P and CTUPEV) and four I(0) variables (PF, E, My and SPA), we will apply the Methodology of Rahbek and Mosconi(1999) introducing the cumulated explanatory I(0) variables into the cointegration relation and

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of Osterwald-Lenum(92). Mackinnon et al.(1999) supplies more correct critical values for the 5 cases of PSS(99).

[18] Among others, see Johansen (1996), Pesaran, Shin and Smith (1999) and Rahbek and Mosconi(1999).

later we will test its exclusion from this relation using the likelihood ratio test. Thus, we will represent the model for study as:

P CTUPEV; csumPF csumE csumMy csumSPA & PF E My SPA

where there are two endogenous I(1) variables (P, CTUPEV) and four exogenous I(1) variables (csumPF, csumE, csumMy, csumSPA) corresponding to the four I(0) variables (PF, E, My, SPA), which are introduced into the short-term model. As we use the variable SPA and not the variable DEF as in equation 1.1, the signal expected in the relation between P and SPA 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 autorregressive 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. The methodology of PSS(99) leads us to choose model IV because we cannot reject the existence of a trend in the long-term relationship at the 10% level (8% to be more accurate). Given VAR(2) and Model IV, one can confirm that the existence of one cointegranting vector cannot be rejected, either by the trace test, or by the maximum eigenvalue test. 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 CTUPEV csumPF csumE csumMy csumSPA t] is given by:[19]

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

where one verifies that the cumulated variables have a relatively high standard error, 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 whose likelihood ratio test follows $\chi^2(5)=4.5391[.475]$. Thus we have:

[19] Between round brackets in the cointegrating vector we have the standard errors.

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

and therefore, the long-term relationship is: P=0.84016 CTUPEV. So, in the long-term, the relationship between inflation rate and the growth rate of unit labour costs is almost unitary.

4.2 Estimation of the Short –Term Model

4.2.1 Initial multi-varied model

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

1) **Equation of ΔP:** (period 1956-1995)

$$\Delta P_t = \underset{[0.276]}{0.94720} - \underset{[0.167]}{0.21085}\Delta P_{t-1} - \underset{[0.154]}{0.13663}\Delta CTUPEV_{t-1} - \underset{[0.341]}{0.075567}PF_{t-1}$$
$$- \underset{[0.074]}{0.25939}E_{t-1} - \underset{[0.097]}{0.12325}My_{t-1} - \underset{[0.457]}{0.21235}SPA_{t-1} - \underset{[0.136]}{0.21484}ECM1_{t-1}$$
$$+ \underset{[0.000]}{0.39909}PF + \underset{[0.000]}{0.43123}E - \underset{[0.918]}{0.0074942}My + \underset{[0.146]}{0.38638}SPA$$

T = 40[1956-1995]; $\overline{R}^2 = 0.69$; SEE = 2.3828; DW = 1.8613;

LM(1, 27) = 0.28330[.599]; RESET(1, 27) = 0.034544[.854]

BJ(2) = 0.16744[.920); HET(1, 38) = 0.73680[.396];

ARCH(2, 26) = 0.89157[.422]

2) **Equation of ΔCTUPEV:** (period 1956-1995)

$$\Delta CTUPEV_t = \underset{[0.991]}{0.014335} + \underset{[0.645]}{0.098067}\Delta P_{t-1} + \underset{[0.163]}{0.18957}\Delta CTUPEV_{t-1} + \underset{[0.622]}{0.055183}PF_{t-1}$$
$$- \underset{[0.003]}{0.64378}E_{t-1} - \underset{[0.095]}{0.19223}My_{t-1} - \underset{[0.007]}{1.1632}SPA_{t-1} + \underset{[0.000]}{1.1673}ECM1_{t-1}$$
$$+ \underset{[0.000]}{0.36720}PF - \underset{[0.315]}{0.14700}E + \underset{[0.908]}{0.011919}My + \underset{[0.084]}{0.65511}SPA$$

T = 40[1956-1995]; $\overline{R}^2 = 0.76$; SEE = 3.3772; DW = 2.0962;

LM(1, 27) = 0.20973[.651]; RESET(1, 27) = 0.6550E-3[.980]

BJ(2) = 0.036334[.982]; HET(1, 38) = 0.27639[.602];

ARCH(2, 23) = 0.81795[.452]

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, and negatively, but only at 10% level, to E_{t-1} and $M_{y\ t-1}$. The long-term relationship (P - 0,84016 CTUPEV) represented by ECM1 presents an expected signal but in this initial model this is not significant even at 10% in contrast to what happens in Rosa(2003).[20] This strengthens the weak exogeneity of the inflation rate in this model.

The positive relation of the variation in inflation with foreign inflation and the variation in the exchange rate corresponds to what would be expected. The negative relation with M_{yt-1} (by the way, almost insignificant) is difficult to explain, but in the parsimonious model, the Wald test suggests its exclusion from the model.

Thus, foreign inflation and the variation in the exchange rate seem to be the main causes of inflation. *Nor the variation in nominal money stock, corrected by the growth rate of real GDP (My), nor the General Government balance in percentage of GDP (SPA) is significant in the equation of* ΔP.

The CTUPEV variation becomes related positively and significantly at 1% to $ECM1_{t-1}$ and PF and negatively at 1% to E_{t-1} and SPA_{t-1}. The explanation for the relation with the first three variables is in Rosa (2003, p. 148). Relative to the negative relation with $SPA_{t-1,}$ a possible explanation could be the fact of a high budget deficit in the previous period implying an increase in the expectations of inflation,[21] which caused wages to increase in the following period.[22] The negative and significant relation at 10% between ΔCTUPEV and My_{t-1} is more difficult to explain, but we do not worry about this, because in the parsimonious equation, the Wald test suggests the exclusion of this variable.

The $ECM1_{t-1}$ is significant at 1% level and close to 1 in the equation of $\Delta CTUPEV$ and it is not significant in the equation of ΔP, so we can conclude that it is the variation in inflation that causes variation in unit labour costs and not the opposite; that is, labour costs seems to respond quickly and significantly to an increase in inflation.

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

[20] However, we must take into account that the ECM1 is slightly different from the one in Rosa (2003)'s model, and therefore not comparable.
[21] The inflationary expectations happened more in a time when the government could use the monetary financing of the deficit.

Relative to the **equation of ΔP**, all the residuals are inside the line bands of double standard deviation[23] and CUSUM and CUSUMSQ tests do not cross any of the significant bars at 5% level.

We estimated the model for period 1956-94 with the aim of leaving an observation for multivaried dynamic forecast. Both the forecasts of ΔP and P, as well as of ΔCTUPEV and CTUPEV for 1995 seem acceptable.

4.2.2 Explicative Parsimonious model of the inflation

We tried to remove from the equation of ΔP in the initial multivaried model the variables that were not significant at the10% 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} and $ECM1_{t-1}$ in the regression of ΔP, despite its non significance in the initial regression. Thus, the Wald test already allows us to reject all the other variables. As E and E_{t-1} have symmetrical coefficients, we substitute them for ΔE and thus the $ECM1_{t-1}$ becomes significant at 10% (equation DPC1 - Table IV of the Annex) and PF_{t-1} becomes significant at 1%; and PF is significant at 1%. Reestimating the previous equation for 1955-88 (equation DPC2), we cannot reject either the predictive capacity after 1988 or the structural stability before and after 1988. CUSUM and CUSUMSQ tests do not indicate problems, moreover.[24]

We tried some dummies,[25] from DPC3 to DPC5 equations, but only dummies Dum87 and Dum80 are significant individually (at 5 and 10 % respectively). The introduction of SME together with Dum87, or with the other two, implies residual autocorrelation (equation DPC5). The Dum87 seems to be the best, always significant at 5% and allows the error-correction mechanism (ECM_{t-1}) to become significant at 5% (see equation DPC3).

In the period 1974-95 without dummies (equation DPC6), the $ECM1_{t-1}$ is not significant, as in the initial model, but the exclusion of this variable (equation DPC7)

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[22] Thus, indirectly, the budget deficit could have a positive influence on inflation through costs, instead of being through demand, as was assumed from the outset in the model.
[23] Better that in the model of Rosa (2003).
[24] Analysis in equation DPC1.
[25] *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

generates autorregressive conditional heteroeskedasticity, so that we opted to keep it. Also, in the period 1974-75, we cannot reject either the predictive capacity after 1988, or structural stability before and after 1988 (equation DPC8).

In 1974-95 the introduction of dummies from equation DPC9 to equation DPC11 allows us to conclude that the Dum87 continues to be significant at 5 % (equation DPC9), but the Dum80 ceases to be significant (equation DPC10) and the introduction of the Dum87, together with SME, does not suffer from autocorrelation (equation DPC11) and the $ECM1_{t-1}$ becomes significant in this last case.

The comparison of the period 1974-95 (equation DPC9) with the period 1955-95 (equation DPC3) allows us to notice a small increase in the absolute value of the coefficients of PF_{t-1}, PF and ΔE in the period 1974-95, to the detriment of the absolute value of $ECM1_{t-1}$.

5. Final 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 effect of the inflation rate on unit labour costs than the reverse. This seems to mean that 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, so we essentially have inflation caused by costs. The strongly significant costs in the short-term relationship are the inflation of imported products (due either to foreign inflation or to the variation in the effective exchange rate).

The comparison of our results with those of other authors allows us to verify that our conclusions are identical to those of the majority of the authors who have made

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conjuncture), *EN* (value 1 up to 1973 - New State), EEC (value 1 after 1986 – Member of the EEC), *SME* (value 1 after 1992 – Participation in the ERM of the EMS).

studies for the 1970s and 1980s, so that one sub-period strongly influences our conclusions. This fact is not strange, because during the New State,[26] the exchange rate of the Escudo was virtually constant and in the 1990s we took measures to control the fluctuation of the same, such as joining the Exchange Rate Mechanism of the European Monetary System. The non-influence of money coincides with the conclusion of Cunha and Machado (1996), but is completely opposed to that of Nunes (1998). However, as we use annual data while Nunes uses quarterly data, and the period is different, any comparison is wrong.[27] Relative to the earliest studies, we must take into account that they do not use the methodology of cointegration, which invalidates the comparison. Santos (1992), concludes that the budget deficit seems to be inflationary, but only in 50% of the analyzed 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,[28] so we therefore do not find our conclusion strange in relation to the non-influence of the budget deficit on the variation in inflation.

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[26] "New State" is the regimen before the April revolution in 1974.

[27] See also Juselius (1992), who finds three significant sources of inflation for Denmark (monetary, wages and imported), while for us, the monetary one is not significant.

[28] There is more evidence therefore that in its model the inflation has contributed towards deficits.

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Annex

Table I - Tests on the existence of two unit roots

Variables	Dickey-Pantula (1987) test					
	1st step				2nd step	
	k	τ_{ρ_2-1}	LM(1) (F version)	Q(4)	k	τ_{ρ_1-1}
U	0	-3.7914[a]	0.7766[.384]	1.667[.797]	0	-1.3184
P	3	-4.3814[a]	3.1880[.084]	1.515[.824]	3	-1.2100
CTUPEV	0	-6.3341[a]	0.7275[.399]	6.510[.164]	0	-2.9512[b]
E	2	-6.0322[a]	0.0028[.958]	0.069[.999]	2	-1.2958
PF	1	-6.6865[a]	1.9630[.170]	4.820[.306]	1	-3.1175[b]
SPA	1	-6.2116[a]	2.9966[.092]	3.511[.476]	1	-1.0536
MY	1	-7.6901[a]	0.7247[.400]	1.458[.834]	1	-1.7117

Notes: Model with a constant; annual data: 1954-1995.
[a] = significant at 1%; [b] = significant at 5%; [c] = significant at 10%.

Table II - Tests on the existence of one unit root: ADF

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)
U	1 (CT)	1	-2.7045	3.7310	-	0.35499	0.65376	0.0229[.881]	0.4167[.981]
	2 (C)	1	-1.3184	-	1.0989	-	0.67161	0.3257[.572]	0.7176[.949]
	3	1	0.0062	-	-	-	-	0.8506[.362]	1.6257[.804]
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]
CTUPEV	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]	-	4.7395[c]	-	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]
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	-	-	-	-		
SPA	1 (CT)	6	-4.0676[b]	8.2772[b]	-	-3.4250[a]	-3.0674[a]	0.0486[.945]	1.5873[.811]
	2 (C)	6	B	-		-			
	3	6	B	-	-	-	-		

Notes: beginning of the tests in models with a trend; annual data: 1954-1995.
[a] = significant at 1%; [b] = significant at 5%; [c] = significant at 10%.
A – we reject the null constant of a time series.
B - we reject the null trend of a time series.

Table IIA - Tests on the existence of one unit root: ADF

Variables	ADF test						
	Mod.	k	$\tau_{\rho-1}$	Φ_1	$\tau_{\mu\mu}$	LM(1) F version	Q(4)
SPA	2 (C)[1]	7	-1.0715	0.7590	-0.60652	1.4605[.239]	1.3502[.853]
	3	7	-0.4856	-	-	1.5424[.226]	1.5606[.816]

Notes: Beginning of the tests in models with a constant, without a trend.
Annual data: 1954-1995.
(1) We begun the selection with k-max=10.

Table III - Tests for a unit root in time series under structural change with endogenous choice of the break point (Tb)

Series	Method	*IO Model*						*AO Model*				
		Tb	k	Estimated Parameters			$t_{\hat{\alpha}=1}$	Tb	k	Estimated Parameters [29]		$t_{\hat{\alpha}=1}$
				$\hat{\theta}$ (DU)	$\hat{\delta}$ (DTb)	$\hat{\alpha}$				$\hat{\theta}$ (DU)	$\hat{\alpha}$	
U	Min $t_{\hat{\alpha}=1}$	1973	1	1.550[a]	-1.047	0.696[a]	-4.50[c]	1972	1	4.012[a]	0.707[a]	-4.19
	Min $t_{\hat{\theta}}$	1985	1	0.009	0.274	0.937[a]	-1.07	1989	1	2.169[b]	0929[a]	-1.56
	Max $t_{\hat{\theta}}$	1973	1	1.550[a]	-1.047	0.696[a]	**-4.50[b]**	1975	1	4.705[a]	0.582[a]	**-3.73[b]**
P	Min $t_{\hat{\alpha}=1}$	1969	5	2.320	-4.014	0.756[a]	-1.61	1983	0	1.338	0.842[a]	-1.99
	Min $t_{\hat{\theta}}$	1983	4	-5.704[a]	6.527	0.974[a]	-0.31	1989	5	-3.007	0.858	-1.44
	Max $t_{\hat{\theta}}$	1969	5	2.320	-4.014	0.756[a]	-1.61	1970	5	12.316[a]	0.859[a]	-0.92
CTUP-EV	Min $t_{\hat{\alpha}=1}$	1971	1	8.054[a]	-5.794	0.265	-4.42[c]	1970	1	10.452[a]	0.265	-4.46[c]
	Min $t_{\hat{\theta}}$	1975	5	-9.246[b]	-14.22[c]	1.453[a]	1.36	1989	0	-0.929	0.683[a]	-2.74
	Max $t_{\hat{\theta}}$	1971	1	8.054[a]	-5.794	0.265	**-4.42[b]**	1972	5	10.943[a]	0.869[b]	-0.39
E	Min $t_{\hat{\alpha}=1}$	1972	1	-4.283[b]	-6.764	0.526[a]	-4.05	1971	1	8.502[a]	0.529[a]	-4.10
	Min $t_{\hat{\theta}}$	1985	3	-3.693[b]	5.466	0.872[a]	-1.09	1988	3	-3.704	0.845[a]	-1.30
	Max $t_{\hat{\theta}}$	1974	1	4.732[b]	-3.915	0.494[a]	-4.03[c]	1975	6	9.753[a]	0.684[a]	-1.60
PF	Min $t_{\hat{\alpha}=1}$	1973	0	-0.269	33.352[a]	0.286[b]	**-6.25[a]**	1973	0	2.464	0.284[b]	**-6.29[a]**
	Min $t_{\hat{\theta}}$	1983	1	-4.877[c]	9.088	0.167	**-4.61[b]**	1984	0	-5.744[b]	0.359[b]	**-4.29[b]**
	Max $t_{\hat{\theta}}$	1970	1	2.818	-2.531	0.218	**-4.32[b]**	1969	1	3.868	0.217	**-4.40[b]**
MY	Min $t_{\hat{\alpha}=1}$	1967	0	8.115[a]	-13.381[b]	0.201	**-5.09[b]**	1967	0	9.802[a]	0.203	**-5.21[a]**
	Min $t_{\hat{\theta}}$	1985	6	-12.299[a]	1.975	0.960[a]	-0.20	1988	1	-2.193	0.629[a]	-2.42
	Max $t_{\hat{\theta}}$	1967	0	8.115[a]	-13.381[b]	0.201	**-5.09[b]**	1968	0	10.556[a]	0.301[b]	**-4.57[a]**
SPA	Min $t_{\hat{\alpha}=1}$	1972	6	-3.756[a]	4.192[b]	0.453[a]	**-4.97[b]**	1974	6	-7.373[a]	-0.111	**-5.34[a]**
	Min $t_{\hat{\theta}}$	1972	6	-3.756[a]	4.192[b]	0.453[a]	**-4.97[b]**	1974	6	-7.373[a]	-0.111	**-5.34[a]**
	Max $t_{\hat{\theta}}$	1961	6	0.714	0.794	0.845[a]	-1.74	1987	3	-2.614	0.909[a]	-1.18

Significance level: [a] = Significant at 1%; [b] = Significant at 5%; [c] = Significant at 10%. **Sample:** 1954-95

Notes: The level of significance refers to the null hypothesis that this coefficient is zero, but for $t_{\hat{\alpha}=1}$ it refers to the null hypothesis of a unit root, according to the Perron and Vogelsang (1992) models. $t_{\hat{\alpha}=1}$ in ***bold*** means that we reject the existence of a unit root, at least at 5 %.

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

AO Model: 1[st] step: $y_t = \mu + \theta DU_t + \tilde{y}_t$

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

[29] However, we put [a], [b] or [c] at $\hat{\alpha}$, there is no mean, because the model has no constant.

Table IV: Parsimonious Equations of ΔP

Dependent Variable: ΔP

Estimation Method: OLS.

ECM1= 1.0000*P -0.84016*CTUPEV

estimated on model: P CTUPEV; csumpf, csume, csummy, csumspa & PF E MY SPA

Equation/ Regressors	DPC1	DPC2	DPC3	DPC4	DPC5	DPC6
	T=41 [55-95]	T_1=34, T_2=7 [55-88]	T=41 [55-95]	T=41 [55-95]	T=41 [55-95]	T=22 [74-95]
Inpt	.092997[.840]	.23312[.650]	.41478[.350]	.36737[.395]	.75194[.100]	-.46639[.550]
PF(-1)	-.23683[.000]	-.25130[.000]	-.28563[.000]	-.26997[.000]	-.28841[.000]	-.25155[.001]
ECM1(-1)	-.12182[.072]	-.12281[.072]	-.14304[.025]	-.12308[.050]	-.14299[.020]	-.083232[.341]
PF	.35997[.000]	.34915[.000]	.37176[.000]	.39077[.000]	.38194[.000]	.39279[.000]
ΔE	.43382[.000]	.48629[.000]	.43141[.000]	.38324[.000]	.37550[.000]	.47112[.000]
Dum80	-	-	-	-4.7750[.086]	-4.5374[.087]	-
Dum87	-	-	-6.6854[.011]	-6.3636[.013]	-7.0791[.005]	-
SME	-	-	-	-	-2.4663[.041]	-
$\bar{R}^2$	.65938	.71523	.70927	.72594	.75174	.73660
SEE	2.4690	2.4326	2.2811	2.2147	2.1079	2.8328
DW	2.4388	2.7737	2.4380	2.3670	2.6857	2.3088
LM$_{(1, T-k-1)}$	2.2668[.141]	6.2749[.018]*	2.1658[.150]	1.5134[.227]	6.6026[.015]*	.56800[.462]
RESET$_{(1, T-k-1)}$	.92664[.342]	.99118[.328]	1.0238[.319]	1.5020[.229]	1.5092[.228]	.52154[.481]
BJ(2)	2.8072[.246]	2.8395[.242]	3.6943[.158]	3.1070[.212]	2.8921[.235]	1.2938[.524]
HET(1, T-2)	.23804[.628]	.34342[.562]	.35303[.556]	.29783[.588]	.41628[.523]	1.3705[.255]
ARCH$_{(2, T-k-2)}$	.55831[.577]	1.4497[.484]	.96366[.392]	1.2862[.290]		2.5660[.110]
Chow$_{(T_2, T_1-k)}$	-	1.1554[.358]	-	-	-	-
Cov$_{(k, T_1+T_2-2k)}$	-	1.5368[.207]	-	-	-	-

Between square brackets: p-value. The null hypothesis is H0: β=0, and is the Student t test for the estimated coefficients.

* Diagnostic test significant at some level indicates the p-value.

Table IV: Parsimonious Equations of ΔP (continuation)

Equation/ Regressors	DPC7	DPC8	DPC9	DPC10	DPC11
	T=22 [74-95]	T_1=15, T_2=7 [74-88]	T=22 [74-95]	T=22 [74-95]	T=22 [74-95]
Inpt	-.84042[.220]	-.57618[.613]	.23662[.754]	.20005[.780]	1.4057[.118]
PF(-1)	-.24943[.001]	-.26189[.004]	-.31259[.000]	-.29712[.000]	-.35700[.000]
ECM1(-1)	-	-.070771[.479]	-.12440[.128]	-.10251[.189]	-.18359[.026]
PF	.39254[.000]	.38648[.000]	.40524[.000]	.42756[.000]	.39262[.000]
ΔE	.51320[.000]	.53826[.001]	.45951[.000]	.40938[.000]	.43000[.000]
Dum80	-	-	-	-4.9465[.113]	-
Dum87	-	-	-6.8446[.031]	-6.5307[.032]	-8.8233[.006]
SME	-	-	-	-	-3.2075[.042]
$\bar{R}^2$	.73721	.79904	.79230	.81365	.83339
SEE	2.8295	2.9698	2.5155	2.3827	2.2530
DW	2.4867	2.9156	2.0011	1.7791	2.6163
LM$_{(1, T-k-1)}$	1.1966[.289]	4.0572[.075]*	.3872E-3[.985]	.18104[.677]	2.8097[.116]
RESET$_{(1, T-k-1)}$	.68751[.419]	.38285[.551]	.91104[.355]	1.9769[.182]	1.3410[.266]
BJ(2)	1.3126[.519]	.29978[.861]	1.2912[.524]	.36885[.832]	.33943[.844]
HET(1, T-2)	2.0063[.172]	2.2945[.154]	.86779[.363]	.30388[.588]	2.0119[.171]
ARCH$_{(2, T-k-2)}$	4.6314[.026]*	.57106[.586]	.23590[.793]		.51311[.610]
Chow$_{(T_2, T_1-k)}$	-	.78102[.618]	-	-	-
Cov$_{(k, T_1+T_2-2k)}$	-	1.1409[.391]	-	-	-

Notes about Table IV

Diagnostic tests: We use the F version of diagnostic tests because Robalo Marques(98) citing Kiviet(86)[30] said that in small samples the F version is preferable. In BJ test we present the LM version following a χ^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)[31]].
RESET – statistic of Ramsey (1969)[32]'s RESET test of functional form misspecification.
BJ – statistic of Jarque-Bera's test of normality of regression residuals [based in Bera and Jarque (1981)[33]].
HET – statistic of Heteroscedasticity test [see Pesaran e Pesaran(1997)]
ARCH – statistic of Autoregressive Conditional Heteroscedasticity test [Engle (1982)[34]'s test]
Chow-statistic of Predictive failure test (2nd test of Chow(1960)).
Cov – statistic of Chow's test of stability of regression coefficients (1st test of Chow(1960)).

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