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# *The P* model and its performance for the Spanish economy*

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The performance of the P* model is tested as an inflation forecaster for the Spanish economy. It is shown that log-run relationships work as expected according to the model and the Quantitative Theory of Money. The Error Correction Model constructed by using the gap between actual prices and the long-term equilibrium price level as an error correction term, offers a consistent explanation for the short-run dynamics in prices. On the other hand, the P* approach shows a forecasting ability similar to that presented for other countries in several studies, although the degree of accuracy in the prediction is not specially satisfactory, mainly for the period 1989:3–1992:3, when the credibility effect generated by the inclusion of the Spanish peseta in the European Monetary System led to an inflation rate much lower than that predicted by the model. Results support the option of a direct inflation target (instead of a monetary aggregate) as the intermediate variable of the monetary policy.

## I. INTRODUCTION

Inflation has been both one of the main problems and a relevant focus of policy action for economic authorities in many countries for the last two decades. In Spain, inflation has always shown worse behaviour than in most developed countries over that time; this is true despite the reduction in the inflation gap between Spain and central countries of the European Union, since that reduction has been slower than desired. Fighting inflation requires a policy to stabilize prices, for which an instrument able to forecast the future path of inflation and to connect it with monetary conditions prevailing in the economy would be very useful.

One of the most interesting proposals in that sense, the P* model, was developed by Hallman, Porter and Small (HPS, hereafter) in 1989, following ideas directly related to a long tradition in mainstream monetary theory (Humphrey, 1989). The model, based on the Quantitative Theory of Money, has generated a broad literature since then, analysing some of the theoretical and econometrical keys of the approach and testing its empirical performance for several countries. A number of papers support the usefulness of this model (HPS, 1989 and 1991; Hallman and Bryden, 1992; Hallman and Anderson, 1993; Deutsche Bundesbank, 1997) while others have relevant criticisms (Christiano, 1989; Funke and Hall, 1994; Hall and Milne, 1994).

In this paper, the performance of the P* model for the Spanish economy is tested by means of a revision of the main technical elements affecting the model and a forecasting exercise implemented for the period 1986–1996. Results show that the model works reasonably well at explaining short-run dynamics for prices and exhibits an ability as a forecaster similar to that for other countries (apart from the period 1989–1992, probably because of a particular credibility effect coming from the inclusion of the peseta in the (at that time) stable EMS – European Monetary System). The average error (around 20% with respect to the actual inflation level) leads to the consideration of this approach as a useful instrument but not a magic solution for monetary authorities. Given these results, the choice of a monetary aggregate as the intermediate target for the monetary strategy of the European Central Bank could lead to an uncertain control of prices.[1]

The plan of the paper is that the next section describes the formulation of the P* model. In Section III, we imple-

[1] Obviously, this remark is only true with respect to the Spanish influence on the overall aggregates.

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ment the P* approach for Spain, with discussion of the problem of the stationarity of money velocity, the identification of the causality direction between prices and money for the Spanish economy, and the presentation of an error correction model to analyse short-term movements in prices. Section IV shows the forecasting exercise for Spanish inflation in the last decade. Finally, Section V draws some general conclusions.

## II. THE P* MODEL

The P* model was developed as an attempt to identify the inflationary potential of a particular economy; in order to do this, the model should be able to forecast future movements in the aggregate price level, giving an evaluation of the inflationary/deflationary character of monetary conditions which prevail in that economy. The P* concept is part of the more general Quantitative Theory of Money, whose basic identity (the 'equation of exchange') is the following:

$$p + y \equiv m + v \quad (1)$$

where $p$ is the aggregate price level, $m$, the money supply, $v$, the velocity of circulation of money and $y$, the real output (all variables are in logarithms).

The model, as developed by HPS (1989, 1991), links the behaviour of the price level to the growth of the money supply by imposing two hypotheses to Equation 1: (i) real output fluctuates around potential real output $(y^*)$; (ii) the velocity has an equilibrium level $(v^*)$, independent of time, that it tracks in the long-term (Hallman and Anderson, 1993). The long-run equilibrium value of the aggregate price level, $p^*$, is then obtained by using $y^*$ and $v^*$:

$$p^* \equiv m + v^* - y^* \quad (2)$$

According to the Quantitative Theory, it is assumed that the actual aggregate price level tends in the long run towards this P* value. From Equations 1 and 2, it is straightforward to show the following relationship among prices, real output and velocities:

$$p - p^* \equiv (v - v^*) + (y^* - y) \quad (3)$$

This implies that if the quantity of money is supporting $p^*$ at a level above $p$, it is depressing the actual velocity below $v^*$ or it is driving real output above $y^*$ (or both). Nevertheless, in the long term, when all adjustments take place and both $v$ and $y$ are on their long-run equilibrium levels, the difference between the current price level and $p^*$ could provide a leading indicator for a future acceleration (if $p^*$ is above $p$) or deceleration (if $p^*$ is below $p$) of inflation.

## III. THE SPANISH ECONOMY AND THE P* MODEL

*Determination of the long-term economy*

As stated in the previous section, the existence of an equilibrium level for the velocity of money towards which the current velocity tends in the long run is one of the basic hypotheses of the P* model. In their seminal work, HPS (1989) considered that the velocity of M2 was stationary for the United States, which implies that shocks on velocity represent just transitory shocks on the level of V2. In contrast, this possibility has generally been discarded in the literature (see Bordes *et al*. 1992; Tödter and Reimers, 1994; Atta-Mensah, 1996).[2]

In Table 1, we show the results from testing the stationarity assumption for the velocity of money in Spain (using ALP, the main monetary aggregate displayed for the Spanish economy).[3] A standard Dickey–Fuller type test, the Phillips–Perron test (1988) was used. Nevertheless, since it has been proved that in finite samples Dickey–Fuller type tests have low power to reject the null hypothesis of a unit root against stable autoregressive alternatives with roots near unity (Kwiatkowski *et al*. 1992), also shown is an analysis based on the so-called KPSS test, which proposes the null hypothesis of stationarity against the alternative of a unit root.

From both tests we obtain that the stationarity of velocity is strongly rejected; hence, the long-run equilibrium for velocity is not the average of the series. The intense and successful process of financial innovation which has happened since the beginning of the 1980s is mainly responsible (Christiano, 1989) for the permanent drop in the value of $v$ detected for several countries. In this context, we need a method to compute $v^*$. Following Hoeller and Poret (1991), the existence of a unit root for velocity leads to an approach in which stochastic tendencies for series are considered and calculated by means of the Hodrick–Prescott detrending method (King and Rebelo, 1993). On the other hand, it is also shown in Table 1 that real output

[2] However, research on this topic was developed by Hallman and Anderson (1993) for M2 in the United States, by implementing several alternative specifications for the evolution of V2; the authors found little evidence against the constant $v^*$ hypothesis.

[3] This aggregate is defined as currency held by the public and sight deposits $(= \mathrm{M1}) +$ savings) deposits $(= \mathrm{M1} + \mathrm{M2}) +$ time deposits, deposits denominated in foreign currency, repo asset sales, asset participations, short-term securities issued by credit institutions, long-term securities issued by deposit money institutions $(= \mathrm{M1} + \mathrm{M2} + \mathrm{M3}) +$ long-term securities issued by official credit institutions and specialized credit institutions, insurance-like liabilities with savings banks, non-interbank private asset transfers, endorsed bills and commercial paper guarantees, treasury notes and bills held by the public and short-term securities of other governments (see Bank of Spain, 1997b). All the series used in this paper come from Bank of Spain (1997a, 1997b).

Table 1. *Unit root and stationarity tests (I)*

| | Phillips–Perron test | | | KPSS Test $(l = 4)$ | |
|---|---|---|---|---|---|
| Variables | $Z(t_{\tilde{\alpha}})$ | $Z(t^*_{\alpha})$ | $Z(t_{\hat{\alpha}})$ | $\eta_\mu$ | $\eta_\tau$ |
| $\Delta y$ | −3.34 * | −3.10 *** | −1.85 * | | |
| $\Delta vALP$ | −8.83 *** | −8.82 *** | −8.04 *** | | |
| $y$ | −2.14 | −1.95 | −0.74 | 2.140 *** | 0.150 *** |
| $vALP$ | −1.90 | −0.85 | 2.72 | 2.000 *** | 0.285 *** |

*Notes*: * and *** denote significance at the 10% and 1% levels, respectively.
$l = \text{int. } \{4(T/100)^{1/4}\} = \text{int. } 4.06 = 4$.
The Phillips–Perron test has been calculated using the long-run variance estimator as proposed by Andrews (1991) and Andrews and Monahan (1992). The critical values are taken from Fuller (1976), Table 8.5.2.c.
The long-run variance has been estimated using the procedure proposed in Newey and West (1987). The critical values are taken from Kwiatkowski *et al.* (1992).

Critical values:

| | 10% | 5% | 1% |
|---|---|---|---|
| $Z(t_{\tilde{\alpha}})$ | −3.15 | −3.45 | −4.04 |
| $Z(t^*_{\alpha})$ | −2.58 | −2.89 | −3.51 |
| $Z(t_{\hat{\alpha}})$ | −1.61 | −1.95 | −2.60 |
| $\eta_\mu$ | 0.347 | 0.463 | 0.739 |
| $\eta_\tau$ | 0.119 | 0.146 | 0.216 |

is an I(1) integrated series for the Spanish economy (as usually happens for most countries) and we make the estimation of potential output $(y^*)$ using the same procedure as for $v^*$. Given these values for the long-run equilibrium level of velocity and output, the aggregate long-run equilibrium for aggregate prices, $p^*$, is obtained from Equation 2.

### *Cointegration and causality*

Having proved that velocity and real output are I(1), we analyse the integration order for the remaining variables (prices and the selected monetary aggregate) which define the exchange equation (Equation 1). We repeat the previous tests and present the results in Table 2.

In these cases, conclusions cannot immediately be drawn since the results for prices and ALP are controversial: whereas according to the KPSS test both series are I(1) (the null hypothesis of stationarity is rejected), using the Phillips–Perron test it is possible to reject the null of a unit root. In order to solve this ambiguity, we follow the suggestions of Perron (1989, 1990) about the possible presence of important structural changes in the trend function. His approach consists of testing for a unit root allowing the possibility of a one-time structural change in that trend function. Here we assume, as did Zivot and Andrews (1992), that the breakpoint is estimated rather than fixed. Three kinds of changes are considered: a change in the intercept (A), a change in the slope (B), or both simultaneously (C) (see Equations 4, 5 and 6, respectively[4]):

$$\text{Model A: } y_t = \mu + \gamma DU_t + \beta t + \delta D(TB)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{4}$$

$$\text{Model B: } y_t = \mu + \beta t + \theta DT^* + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{5}$$

$$\text{Model C: } y_t = \mu + \gamma DU_t + \beta t + \theta DT^* + \delta D(TB) + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{6}$$

Additionally, two kinds of model are tested. The first one, called the Additive Outlier Model (AOM) considers the change to be happening instantaneously and not affected by the dynamics of the series. The second one, called the Innovational Outlier Model (IOM), assumes that the possible break occurs slowly over time.[5]

[4] The different dummies introduced in these models have the following implications: $D(TB)$ tries to pick up the presence of an anomalous year in the series; $DU$ tries to capture a change in the intercept, while the change in the slope is expected to be detected by $DT^*$. The variable of interest is always $y$; $D(TB) = 1$ if $t = TB$; 0 otherwise; $DU = 1$ if $t > TB$, 0 otherwise and $DT^* = t - TB$ if $t > TB$, 0 otherwise. $TB$ represents the breakpoint.

[5] For a complete presentation of the approach we are using in this part of this paper, see specially Perron (1989, 1990), Perron and Vogelsang (1992), Zivot and Andrews (1992) and Vogelsang and Perron (1994).

Table 2. *Unit root and stationarity tests (II)*

| Variables | Phillips–Perron test | | | KPSS Test $(l=4)$ | |
|---|---|---|---|---|---|
| | $Z(t_{\tilde{\alpha}})$ | $Z(t_{\alpha}^{*})$ | $Z(t_{\hat{\alpha}})$ | $\eta_{\mu}$ | $\eta_{\tau}$ |
| $\Delta p$ | −3.67 ** | −2.09 | −0.89 | | |
| $\Delta ALP$ | −8.73 *** | −5.92 *** | −1.70 | | |
| $p$ | −1.21 | −3.61 ** | 6.88 | 2.175 *** | 0.544 *** |
| $ALP$ | −1.40 | −4.13 ** | 7.03 | 2.177 *** | 0.552 *** |

*Notes*: ** and *** denote significance at the 5% and 1% levels, respectively.
$l = \text{int.}\ \{4(T/100)^{1/4}\} = \text{int.}\ 4.06 = 4$.
The Phillips–Perron test has been calculated using the long-run variance estimator as proposed by Andrews (1991) and Andrews and Monahan (1992). The critical values are taken from Fuller (1976), Table 8.5.2.c.
The long-run variance has been estimated using the procedure proposed in Newey and West (1987). The critical values are taken from Kwiatkowski *et al.* (1992).

Critical values:

| | 10% | 5% | 1% |
|---|---|---|---|
| $Z(t_{\tilde{\alpha}})$ | −3.15 | −3.45 | −4.04 |
| $Z(t_{\alpha}^{*})$ | −2.58 | −2.89 | −3.51 |
| $Z(t_{\hat{\alpha}})$ | −1.61 | −1.95 | −2.60 |
| $\eta_{\mu}$ | 0.347 | 0.463 | 0.739 |
| $\eta_{\tau}$ | 0.119 | 0.146 | 0.216 |

Table 3. *Unit root tests allowing for structural breaks*

| Series | Model | $TB$ | $K$ | $\hat{\alpha}$ | $t\hat{\alpha}$ |
|---|---|---|---|---|---|
| P | AOM-A | 90:4 | 4 | 0.98 | −2.04 |
| P | AOM-B | 84:1 | 4 | 0.94 | −3.88 |
| P | AOM-C | 84:1 | 4 | 0.94 | −4.00 |
| P | IOM-A | 75:4 | 5 | 0.98 | −2.54 |
| P | IOM-C | 76:1 | 5 | 0.97 | −4.63 |
| ALP | AOM-A | 94:4 | 4 | 0.98 | −1.96 |
| ALP | AOM-B | 91:3 | 4 | 0.98 | −3.36 |
| ALP | AOM-C | 91:2 | 4 | 0.97 | −1.84 |
| ALP | IOM-A | 75:1 | 4 | 0.98 | −2.26 |
| ALP | IOM-C | 77:4 | 5 | 0.96 | −3.93 |

$TB$ = breakpoint.
$K$ = truncation lag parameter.

Critical values:

| Model | 10% | 5% | 1% |
|---|---|---|---|
| AOM-A | −4.79 | −5.06 | −5.68 |
| AOM-B | −4.38 | −4.65 | −5.28 |
| AOM-C | −5.11 | −5.40 | −5.96 |
| IOM-A | −4.82 | −5.10 | −5.70 |
| IOM-C | −5.25 | −5.55 | −6.21 |

A summary of the results obtained using this approach is given in Table 3. The main conclusion is that for every series under analysis, the null hypothesis of the existence of a unit root after allowing for a break is not rejected. Hence, it has been established that all variables in the equation of exchange are integrated of order one for the Spanish case.

The validity of the P* model as a good representation of the data requires that $p$ and $p^*$ form a cointegrating pair in the sense of Engle and Granger (1987). In the next subsection we present this condition in an alternative way and connect it with the use of an error correction model to explain movements in prices. Now, however, the third line of Table 4 shows the result from the Phillips–Hansen

Table 4. *Cointegration relationships*

| Variables | Phillips–Ouliaris test $\hat{Z}_\alpha$ ($H_0$: non-cointegration) |
|---|---|
| $(P, P^* ALP)$ | $-16.54$* |
| $(P, ALP)$ | $-2.76$ |

*Notes*: * denotes significance at the 10% level.
Method by Phillips and Hansen (1990) and Hansen (1992), based on the procedure suggested by Andrews (1991).
The critical values come from Haug (1992) Table 2. These critical values are: $-16.24$ for the 10% significance level, $-19.29$ for the 5% and $-26.18$ for the 1% (with $T = 100$) and for a deterministic cointegration (no linear trend in the long-run regression).
$P^* ALP$ means the value for $P^*$ by using $ALP$ as the monetary aggregate in the Equation (2).

procedure to estimate cointegration relationships. This method starts (Phillips and Hansen, 1990) with the OLS (ordinary least squares) estimates obtained from the static regression of Engle–Granger; after that, two non-parametric corrections are introduced to eliminate possible serial correlations in residuals and the presence of endogeneity of regressors. The validity of the P* approach at this level is supported by the rejection of the null hypothesis of non-cointegration between the actual level of prices and the long-run equilibrium level of aggregate prices. On the other hand, we study the possible cointegration relationship between prices and the ALP aggregate. The fourth line of Table 4 shows that, following the same procedure described above, it is not possible to reject the non-cointegration hypothesis. This leads to the conclusion that money does not solely determine prices in the long run for the Spanish economy.

Hall and Milne (1994) and Funke and Hall (1994) emphasized the fact that the validity of the P* approach to represent the data is one thing and the implicit assumption of the model assuming that money causes prices is quite another. A causality analysis should reveal if this assumption is true rather than the counterview that prices cause money.

We study the causality problem by following Granger (1969 and 1988)[6], and test for the null hypothesis of non-causality from $Y$ to $X$ in the following expression:

Table 5. *Results from the causality analysis*

| Variable caused ($X$) | Variable causing ($Y$) | Lags for $X$ | Lags for $Y$ | F-stat. |
|---|---|---|---|---|
| $ALP$ | $P$ | 8 | 2 | 0.14 |
| $P$ | $ALP$ | 12 | 2 | 6.02 *** |

*Notes*: Test of Granger with null hypothesis of non-causality.
*** denotes significance at the 1% level.
Critical values for F with (2, 80) degrees of freedom (both cases in the table) are 3.13 (5%) and 4.92 (1%).
The optimal number of lags have been calculated following the Final Prediction Error (FPE) procedure presented in Hansen (1989). The optimal number of lags for both lagged variables in Equation 8 must minimize

$$FPE = \frac{T+q}{T-q} \times \frac{1}{T} \times SSR$$

where $T$ is the number of data points, $q$ the number of regressors and $SSR$ the squared sum of residuals.

$$\Delta X_t = \alpha_0 + \sum_{j=1}^{m} \alpha_{1j} \Delta X_{t-j} + \sum_{j=1}^{n} \alpha_{2j} \Delta Y_{t-j} + \varepsilon_{1t} \qquad (7)$$

(an equivalent expression should be tested for $Y$ not being caused by $X$).

Results from this exercise are shown in Table 5. Unlike the conclusions for other economies (Funke and Hall, 1994; Hall and Milne, 1994; Hansen and Kim, 1996), it seems that causality between money and prices works precisely in the direction expected by the P* approach, given the rejection of the null hypothesis of non-causality from the monetary aggregate to prices for Spain. Nevertheless, caution is necessary in drawing conclusions from these causality tests since they are quite often very sensitive with respect to the choice of the optimal lag length (Hansen, 1989).

*Short-run dynamics*

The analysis of short-run movements in prices through an Error Correction Model (ECM) structure, where the inflation rate is explained by its own lagged values and by the gap between the actual value of the overall price level and

[6] If both I(1) series (in this case prices and money) were cointegrated, it would be necessary to improve the specification of the test by including the term expressing that long-run relationship (see, for instance, Corradi *et al.* 1990). In such a case, both long-run and short-run causalities can be analysed. We can formulate it as follows:

$$\Delta X_t = \alpha_0 + \gamma_1 (X_{t-1} - \varphi Y_{t-1}) + \sum_{j=1}^{m} \alpha_{1j} \Delta X_{t-j} + \sum_{j=1}^{n} \alpha_{2i} \Delta Y_{t-i} + \varepsilon_{1t}$$

A similar representation would be stated for $Y_t$. It has to be tested that:

(1) $H_0$: $\gamma_1 = 0$ (non-causality in the long run)
(2) $H_0$: $\alpha_{2i} = 0$ (non-causality in the short run) for any $i$, $i = 2, \ldots, n$.

Table 6. *Stationarity of the P* gap*

| Variables | Phillips–Perron test | | | KPSS Test $(l=4)$ | |
|---|---|---|---|---|---|
| | $Z(t_{\tilde{\alpha}})$ | $Z(t_{\alpha}^{*})$ | $Z(t_{\hat{\alpha}})$ | $\eta_{\mu}$ | $\eta_{\tau}$ |
| *P-P *ALP* | −3.67 ** | −2.09 | −0.89 | | |

*Notes*: **denotes significance at the 5% level.
$l = \text{int.}\ \{4(T/100)^{1/4}\} = \text{int.}\ 4.06 = 4$.
The Phillips–Perron test has been calculated using the long-run variance estimator as proposed by Andrews (1991) and Andrews and Monahan (1992). The critical values are taken from Fuller (1976), Table 8.5.2.c.
The long-run variance has been estimated using the procedure proposed in Newey and West (1987). The critical values are taken from Kwiatkowski *et al.* (1992).

Critical values:

| | 10% | 5% | 1% |
|---|---|---|---|
| $Z(t_{\tilde{\alpha}})$ | −3.15 | −3.45 | −4.04 |
| $Z(t_{\alpha}^{*})$ | −2.58 | −2.89 | −3.51 |
| $Z(t_{\hat{\alpha}})$ | −1.61 | −1.95 | −2.60 |
| $\eta_{\mu}$ | 0.347 | 0.463 | 0.739 |
| $\eta_{\tau}$ | 0.119 | 0.146 | 0.216 |

its long term equilibrium value (P* gap, hereafter), this second component being the error correction term (as proposed by HPS[7]) requires that in the long run this gap vanishes. In other words, as pointed out by Allen and Hall (1990), for such a model to be valid, the difference between $p$ and $p^{*}$ must be stationary; if we write

$$\Delta p_t = \alpha_0 + \alpha_1 (p_{t-1} - p_{t-1}^{*}) + \sum_{j=1}^{n} \beta_j \Delta p_{t-j} \qquad (8)$$

as a consequence of $\Delta p_t$ being I(0) (see the previous subsection), which implies $\Delta p_{t-j}$ is I(0) for any value of $j$, it is necessary that the error correction term also be stationary to have a consistent equation. In Table 6, the results offered by the Phillips–Perron and KPSS tests clearly support the hypothesis of stationarity for the P* gap.

Given these results, the ECM is estimated and the results (jointly with the specification tests) are shown below. As expected according to P* model principles, the coefficient $\alpha_1$, which corresponds to the lag between actual prices and their long-run equilibrium level, is highly significant and negative: if prices are higher than that equilibrium value, it will cause a slowdown in inflation. As a whole, the results show that short-run price movements for the Spanish economy are reasonably well described by considering their intrinsic movements and the $p^{*}$ gap.

Short term equation (1970:1–1996:3):

$$\Delta p_t = \alpha_0 + \alpha_1 (p_{t-1} - p_{t-1}^{*}) + \sum_{j=1}^{n} \beta_j \Delta p_{t-j}$$

| Coefficient | Value | $t$ |
|---|---|---|
| $\alpha_0$ | 0.01 | 0.35 |
| $\alpha_1$ | −0.12 | −4.88 |
| $\beta_1$ | 0.97 | 10.89 |
| $\beta_2$ | −0.32 | −3.33 |
| $\beta_4$ | 0.42 | 4.81 |
| $\beta_5$ | −0.34 | −2.98 |
| $\beta_6$ | 0.25 | 3.86 |

R-squared = 0.88
DW = 2.03
Heteroscedasticity (Goldefeld–Quandt test): F $(42,35) = 0.319$ (no rejection of homoscedasticity)
ARCH $\{2\}$: Chi-squared $(2) = 0.12$ (no rejection of time-variable homoscedasticity)
Residual correlation (General LM test): Chi-squared $(2) = 1.64$ (no rejection of no serial correlation).

## IV. FORECASTING PERFORMANCE OF THE P* MODEL

The final evaluation for the P* model must necessarily rest in an analysis of its ability to do what it is supposed to do: forecasting. One of the most critical elements for policy-makers is precisely the need to forecast the future behaviour of relevant macroeconomic variables. This kind of exercise is difficult and, although the P* model seems to be consistent with the widely accepted Quantitative Theory of Money (given the expected direction in causality relationships), this consistency means nothing in terms of P*'s ability to predict inflation in the short-term, since the

[7] In their seminal papers in 1989 and 1991, HPS formulated an ECM on acceleration of inflation; although it is not explicitly stated, it must be probably due to the fact that only the second difference of prices is stationary for the US data.

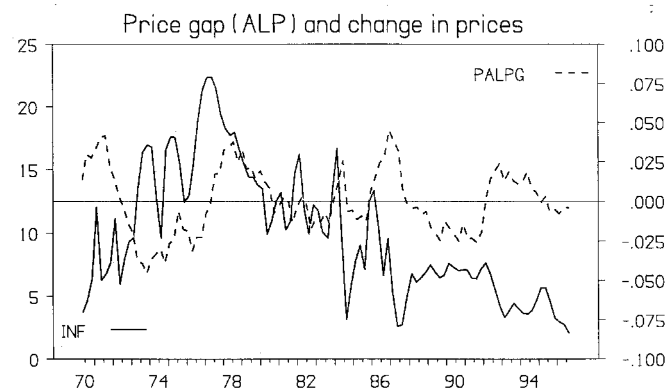


Fig. 1. *The P* gap and the rate of inflation for the Spanish economy*

Quantitative Theory is vinculated with the long run (Christiano, 1989). However, only by testing empirically its performance as a leading indicator for inflation, can the P* approach be considered as a useful (or useless) tool for monetary authorities.

A first qualitative approximation for the Spanish case may be obtained from a graphic representation of the P* gap and the actual inflation (see Fig. 1). Being defined as $(p - p^*)$, when that gap is negative (positive), the actual price level is below (above) the long-term equilibrium price level, which implies that prices should exhibit an upward (downward) path. Generally speaking, it is possible to find those features in the figure (quite clearly, for instance, from 1973 to 1977, from 1978 to 1980 and at the end of the overall period). However, at the end of the 1980s and the beginning of the 1990s, the Spanish inflation rate remained stable despite the expected increase, given the negative $p^*$ gap. We will come back to this special subperiod later.

A more rigorous analysis is made by implementing a rolling forecast. The short-term equation (8) is estimated over a sample of fixed length;[8] both one-year-ahead and three-year-ahead forecasts are computed. Then, the next observation in the sample is added and the last one dropped, and this process is repeated for each term for the 1986:1–1996:3 period. The results are given in Table 7, with the forecasting for the first and the third terms in each year (in annual rates) and several summary statistics.

Three different periods are taken into account to analyse these results: the overall sample, the subsample 1989:3–1992:3 and the remaining subsample. The singularity of that particular three-year period comes from the fact that it was the time from the incorporation of the Spanish currency to the European Monetary System (19/06/1989) to the huge crisis in the EMS after Denmark refused the Maastricht Treaty as the basis for the Economic and Monetary Union (02/06/1992). The crucial fact which can be drawn from Table 7 is the large loss of validity of the model for that particular subperiod. For all proposed statistics, there is an increase in the size of the measurement error with respect to the rest of the period for the one-year-ahead and the three-year-ahead forecasts. By focusing on the simple mean error it is interesting to note that, whereas for the remaining period, the P* approach underpredicts the actual inflation, between 1989:3 and 1992:3 it overpredicts prices growth (by so much that this effect dominates the results for the whole period). The other two measures give the proof of the worst behaviour of the model for those special years.[9] Before developing an economic explanation of this singularity we will make a brief comparison with other authors' results in order to state the relative degree of accuracy of the forecasting exercise for the Spanish case. Table 8 shows the comparison.

In Table 8, we construct an indicator, the mean absolute error relative to the mean level of actual inflation, which is much more useful to compare research in different countries.[10] General results[11] seem to show worse behaviour of the model for Spain (a 27% deviation) than for other cases (between 17% and a 20% deviation). Nevertheless, if we exclude the key period 1989:3–1992:3, which presents a high 38% deviation, the degree of accuracy is approximately the same for the Spanish economy, a deviation of 20%. Hence, we can conclude that, generally speaking, the P* approach has a reasonable prediction ability, although, at a time when the concern about inflation is so extreme (and good results have been achieved in most developed countries in the anti-inflationary fight), a 20% mistake in the prediction exercise is not irrelevant. This kind of result for a particular country offers a guide to the choice of monetary framework by the European Central Bank. The discussion about using an inflation target or a monetary aggregate as the intermediate variable is crucially based on the stability of the link between the selected monetary aggregate and prices.[12] If that stability exists, a monetary aggregate could play the role of the optimal

[8] The first sample is for the 1970:1–1985:4 period (60 data points).

[9] Although in some work (Bordes *et al.* 1989), simple mean is used to draw conclusions on the validity of forecasting results, it is quite obvious that, by means of such a rough measure, some mistakes (if they are of opposite sign and compensate each other) would lead to an apparent, and false, greater valuation of the approach.

[10] If we directly took the mean absolute error, a mistake of one point would always be worse than a 0.5 points error, which is not true if, for instance, the first country has inflation as much as three times that of the second country.

[11] These results are for the one-year-ahead forecast and those indicators which do not exist in the quoted research are from our own elaboration.

[12] A large number of analyses have been made for the last years on this topic. See Begg (1997) and Ramaswamy (1997) as two relevant and recent examples.

Table 7. *Results from the forecast exercise (1986:1–1996:3)*

| Term | Actual inflation (1) | Forecast inflation (1 year ahead) (2) | Gap (2)–(1) | Forecast inflation (3 years ahead) (4) | Gap (4)–(1) |
|---|---|---|---|---|---|
| 86:1 | 13.3 | 12.7 | −0.6 | 12.8 | −0.5 |
| 86:3 | 6.6 | 6.0 | −0.6 | 8.2 | 1.6 |
| 87:1 | 5.2 | 5.1 | −0.1 | 6.1 | 0.9 |
| 87:3 | 6.7 | 6.6 | −0.1 | 6.9 | 0.2 |
| 88:1 | 6.8 | 6.8 | 0 | 4.0 | −2.8 |
| 88:3 | 6.4 | 5.0 | −1.4 | 3.5 | −2.9 |
| 89:1 | 7.4 | 5.4 | −2.0 | 4.7 | −2.7 |
| 89:3 | 6.4 | 7.9 | 1.5 | 6.6 | 0.2 |
| 90:1 | 7.6 | 9.8 | 2.2 | 9.3 | 1.7 |
| 90:3 | 7.0 | 10.7 | 3.7 | 8.9 | 1.9 |
| 91:1 | 7.0 | 11.0 | 4.0 | 11.0 | 4.0 |
| 91:3 | 6.4 | 10.3 | 3.9 | 11.6 | 5.2 |
| 92:1 | 7.6 | 9.8 | 2.2 | 11.2 | 3.6 |
| 92:3 | 5.5 | 6.5 | 1.0 | 8.8 | 3.3 |
| 93:1 | 3.3 | 4.4 | 1.1 | 6.4 | 3.1 |
| 93:3 | 4.4 | 3.5 | −0.9 | 5.3 | 0.9 |
| 94:1 | 3.6 | 2.9 | −0.7 | 4.2 | 0.6 |
| 94:3 | 3.9 | 1.7 | −1.8 | 2.2 | −1.7 |
| 95:1 | 5.6 | 1.9 | −3.7 | 2.0 | −3.6 |
| 95:3 | 4.5 | 2.7 | −1.8 | 1.8 | −2.7 |
| 96:1 | 3.0 | 4.0 | 1.0 | 2.3 | −0.7 |
| 96:3 | 2.0 | 3.0 | 1.0 | 2.0 | 0 |

| Period | Forecasting | Mean error | Mean absolute error | Sq. root of mean squared error | Mean abs. error/mean level of inflation |
|---|---|---|---|---|---|
| Overall | 1 year ahead | 0.36 | 1.60 | 2.01 | 27% |
| Overall | 3 years ahead | 0.44 | 2.04 | 2.48 | 34% |
| 1989:3–1992:3 | 1 year ahead | 2.64 | 2.64 | 2.87 | 38% |
| 1989:–1992:3 | 3 years ahead | 2.86 | 2.86 | 3.24 | 42% |
| Rest | 1 year ahead | −0.70 | 1.12 | 1.45 | 20% |
| Rest | 3 years ahead | −0.96 | 1.65 | 2.03 | 30% |

*Notes*: Results expressed in annual rates.
The mean averages of actual inflation for Spain over the periods were 5.90 (the overall period), 6.80 (1989:3–1992:2) and 5.50 (the rest of the period).

intermediate target for the monetary policy of the European Central Bank. The results from the different studies for single European countries (or groups of them) show a large disparity.[13] Our analysis strengthens the alternative of a direct inflation target.

The remaining and quite interesting fact to be explained here is the basis for the serious deficiency that the model shows in its forecasting ability between 1989:3 and 1992:3 (five out of the six highest errors for a singular term are within this short subperiod, and the model overpredicts by more than two and a half points the actual inflation). The reason for the high inflation rate forecast by the P* approach is the strong rhythm of growth of monetary aggregates at the end of the 1980s and the beginning of the 1990s in Spain, even above the upper limit of the target stated by the Bank of Spain (Ayuso and Escrivá, 1997). Several factors helped to generate that path: first, an expansive stage of the economy; second, a procyclical fiscal

[13] See Browne *et al.* (1997) for a highly detailed survey of many of those studies.

Table 8. *Comparison among forecasting results for several countries*

| Research | Mean error | Mean abs. error (2) | Mean level of actual inflation (3) | (2)/(3) |
|---|---|---|---|---|
| HPS (1991)– U.S. (1971–1989) | −0.19 | 1.03 | 5.97 | 17% |
| Bordes *et al.* (1992)– France (1989:1–1990:4) | 0.32 | 0.47 | 2.78 | 17% |
| Bordes *et al.* (1992)– Germany (1988:3–1990:2) | 0.52 | 0.52 | 2.56 | 20% |
| * Spain (1986:1–1996:3) | 0.36 | 1.60 | 5.90 | 27% |
| * Spain (1989:3–1992:3) | 2.64 | 2.64 | 6.80 | 38% |
| * Spain (1986:1–1989:2/1992:4–1996:3) | −0.70 | 1.12 | 5.50 | 20% |

*Note*: * These results come from our research (see also in Table 7).

policy (not only but specially because of the need for expanding the Spanish Welfare State to approximate it to European Union standards) which additionally supported the growth; and, last but not least, the peseta entered into the EMS with a very strong (it would not be very risky to write an overvalued) central exchange rate. The necessity of fighting inflation and of maintaining the official exchange rate drove monetary authorities to generate an increasing interest rate differential between Spain and most of OECD countries; given the progressive liberalization of foreign capital movements which concluded in 1992, a huge quantity of short-run speculative capital (jointly with more stable investment caused by the favourable economic situation) came into Spain and reinforced monetary growth. Nevertheless, there were a couple of elements which explain why the inflation rate was lower than the P* approach predicts given monetary aggregates path. The first one was the positive impact on inflationary expectations (in the sense of reducing them) of the credibility gained because of membership of the Spanish currency in the EMS, whose stability was almost absolute at that time. In addition to this estremely relevant reputational effect,[14] the 'hard peseta' resulted in lower costs of Spanish imports.

The EMS crisis, which happened from the second half of 1992 after the Danish disapproval and the French reluctant acceptance (in their respective referenda) of the Maastricht Treaty, was due to the very obvious incompatibilities between economic fundamentals and the exchange rate for some countries belonging to the System (Spain among them), suddenly eliminated those two positive elements, with four consecutive devaluations of the Spanish currency. Hence, from 1993 to 1996, the P* approach makes up for its lost forecasting ability and even underpredicts actual inflation (as before 1989). This underprediction out of such a particular stage (1989:3–1992:3) probably is at least partially vinculated to the factor which has been quite often (both in academic analysis and in the mass-media) blamed for the Spanish higher inflation (with respect to OECD countries): structural rigidities in the Spanish economy (excess of regulations in the labour market, lack of competition in many service activities, public monopolies and so on).

## V. CONCLUDING REMARKS

In this research, the performance of the P* model as an inflation forecaster for the Spanish economy has been tested. It has been shown that (after proving that the velocity of money is not stationary for Spain, regardless of the monetary aggregate under consideration) long-run relationships work as expected according to the model and the Quantitative Theory of Money (although money does not solely determine prices in the long run); the error correction model constructed by using the gap between actual prices and the long-term equilibrium price level as the error correction term, offers a consistent explanation for the short-run dynamics in prices. On the other hand, the P* approach shows a forecasting ability similar to that presented for other countries in several papers. It is not true for the period 1989:3-1992:3, when the credibility effect generated by the inclusion of the peseta in the EMS led to an inflation rate much lower than that predicted by the model (which does not mean it was low in absolute terms or in relative terms with respect to other developed countries). Apart from that time, the average deviation in relation with the mean level of actual inflation is about 20% (sometimes overpredicting but generally underpredicting the change in prices, which may be due to the impact of structural rigidities in the Spanish economy which are not considered by this approach). Although it is possible to state that the P* model can be a useful instrument for the implementation of monetary policies, our results show that, following Christiano (1989), 'the P*

[14] Ayuso and Escrivá (1997) offer a clear development of the importance for Spanish monetary policy of the credibility induced by the inclusion of the peseta in the EMS.

model is not the inflation forecaster's Holy Grail'. According to our results, the choice of an inflation target as the intermediate variable of the monetary policy could be a good alternative to the choice of a monetary aggregate, for the Spanish economy.

Finally, an interesting extension of this research could be the introduction of several improvements in the error correction model specification for the short-run dynamics in price analysis. Additional variables such as commodity prices (Hallman and Bryden, 1992), import prices (Bundesbank, 1992; Tödter and Reimers, 1994) and a dummy variable (HPS, 1989, with a dummy for the oil shock in 1973; in our case, it should try to pick up the credibility effect of the EMS) could increase the explicative ability of the ECM.

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