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Banco de España - Servicio de Estudios
Documento de Trabajo nº 9422

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(*) This paper has benefited from the comments and suggestions by J Ayuso, J. J. Dolado, F. Restoy, C. Sánchez, T. Sastre and J. Vallés.

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ISBN: 84-7793-363-4
Depósito legal: M-2359-1995
Imprenta del Banco de España

ABSTRACT

Using several recent tests for structural stability in regressions with I(1) variables and for the existence of cointegration in models with regime shifts, the empirical evidence available on the existence of a structural break in the Spanish long-run demand for broad money (ALP) is analysed.

Results indicate that shifts affecting the demand for ALP in recent years have substantially altered its long-run properties. As to the cause of this structural break, emphasis is placed on the role played by the increasing openness of the Spanish financial system to international markets as obstacles to free capital movements have progressively disappeared.

1. INTRODUCTION

In 1978 the Banco de España announced, for the first time, monetary targets in the form of an annual growth range for M3. The announcement of targets finalised a lengthy process, initiated in 1973, to design a monetary policy framework that had the course of a broad monetary aggregate as its intermediate target. From the outset, the existence of a stable relationship between the final variables and the monetary magnitudes was -as in other countries- the essential criterion for selecting the above-mentioned target.

During the eighties, the basic monetary policy framework was subject to numerous alterations. An instrumental variable based on interest rates, rather than on bank reserves, was adopted. The definition of the intermediate target was extended to become ALP (liquid assets held by the public) as from 1984. The ALP aggregate was also modified on successive occasions to accommodate financial innovations, which increased the degree of substitutability between instruments included and not included in its definition. The eighties also saw an important conceptual change in the monitoring of money supply growth targets. As the difficulties of retaining a strict short-term control of liquidity progressively surfaced, and as the exchange rate of the peseta became part of the Banco de España's reaction function, the monetary control strategy acquired a more gradualist bent whereunder short-term deviations in targets were considered to be not very informative.

The entry of the peseta into the ERM in June 1989 was undoubtedly a landmark in this process. With exchange rate fluctuation bands of ±6% vis-à-vis the central parities, the behaviour of the monetary aggregates became -more than a target in itself- an economic policy reference that warned of possible inconsistencies between domestic and external targets. Nonetheless, the capacity of the authorities to respond to undesired trends in monetary aggregates was in practice limited by the foreign exchange constraint.

The widening of the fluctuation bands of ERM currencies in August 1993 provided a new ingredient for the design of monetary policy, opening up the possibility that, in return for the diminished nominal anchorage of the exchange rate, greater attention might be paid to developments in certain domestic indicators. From this standpoint, a broad monetary aggregate such as ALP was an obvious candidate to play a prominent role. However, the view is now widely held that the historical relationships, which once made ALP the monetary aggregate whose performance the Banco de España chose as its prime focus of attention, have deteriorated considerably[1]. This view is based not only on the signals sent by the aggregate's shorter-term behaviour, but also on the longer-term relationship itself, suggesting the existence of a significant structural break in the demand for ALP.

With this perspective, the paper analyses the available empirical evidence on the existence of a structural break in the ALP long-run demand function that would invalidate the estimated historical relationships. Also, several hypotheses on the nature of this structural change are tested. To this end, section 2 looks at various recently proposed tests for structural stability in regressions with first-order integrated variables and for the existence of cointegration in models with regime shifts. Section 3 draws the main conclusions.

2. LONG-RUN DEMAND FOR ALP: COINTEGRATION AND STABILITY

Since the work by Dolado (1988), empirical studies on the estimation of money demand functions in Spain have -in congruence with international evidence in this respect- stressed first, the existence of a long-run equilibrium relationship between integrated variables

[1] In this connection see, for example, the Banco de España 1993 Annual Report.

(cointegration), interpretable as a real-balances demand function; and further, the estimation of the dynamic process of adjustment to this equilibrium in the form of an error correction mechanism (ECM) model.

Thus, in the latest paper published on this subject (Cabrero, Escrivá and Sastre, 1992), demand equations are estimated for several monetary aggregates as follows:

$$\Delta(m-p)_t = k - \alpha[m-p-m^d(.)]_{t-1} + \delta_0(L)\Delta(m-p)_{t-1} + \delta_1(L)\Delta y_t + \delta_2(L)\Delta r^p_t + \delta_3(L)\Delta r^a_t + \delta_4(L)\Delta^2 p_t + \epsilon_t \quad (1)$$

$$m^d_t(.) = \alpha_0 + \alpha_1 y_t + \alpha_2 r^p_t + \alpha_3 r^a_t + \alpha_4 \pi_t \quad (2)$$

where y_t, p_t, r^p_t, r^a_t and π_t are[2], respectively, real GDP, the consumer-price index (CPI), the weighted average rate -net of taxes- of the assets included in the monetary aggregate, an interest rate representative of alternative assets (internal rate of return -net of taxes- on government debt) and the rate of inflation measured by the CPI[3]; Δ is the difference operator and $\delta_i(L)$ are finite-order polynomials in the lag operator.

In the case of the enlarged monetary aggregate ALP2 (ALP plus commercial paper), the existence of cointegration in the period 1979/I-1989/II was determined, while at the same time it was assessed

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[2] See Cabrero et al. (1992) for an accurate definition of the variables and their sources.

[3] The order of integration of the variables in equation (2) has been studied in many papers: Cabrero et al. (1992), Ayuso and Vega (1994), etc. Generally, it holds that the monetary aggregates and prices are I(2), while the monetary aggregates in real terms, income, inflation and interest rates are I(1).

that certain short-run circumstantial instabilities were co-existing with a high degree of stability of the long-run relationship, although no formal tests were performed in this latter respect. Other authors -Escrivá and Malo de Molina (1991) and Dolado and Escrivá (1992)- have highlighted the same type of findings.

Given the emphasis placed on the long-run properties of the monetary aggregate, this paper focuses on the long-run demand for real balances relationship given by equation (2) and attempts to analyse whether the passage of time makes it necessary to alter any of the foregoing conclusions and, if so, in which sense. From this standpoint, the parameters describing the short-run dynamics are treated here as nuisance parameters on which we shall not focus. One of the paper's aims is also the application of estimation methods which, beyond the traditional result of superconsistency (Stock, 1987), allow for the use of standard (asymptotic) inference procedures on the long-run parameters.

Table 1 shows equation (2) estimates over the entire sample 1979/I-1993/IV and several residual-based tests that help evaluate the empirical evidence on the existence or not of cointegration in the period analysed. Estimation is performed using Phillips and Hansen (1990) fully modified (FME) procedure and the long-run variance-covariance matrix is estimated, following suggestions by Andrews (1991) and Andrews and Monahan (1992), using a prewhitened quadratic spectral kernel with a first-order autoregression for the prewhitening and an automatic bandwidth estimator. The FME is asymptotically equivalent to maximum-likelihood and provides, on the one hand, for the mitigation of second-order biases detected in static OLS estimates of cointegrating vectors, and on the other, for the use of standard (asymptotic) procedures for inference.

Table 1

Estimation method: FME;	**Sample: 1979/I-1993/IV**

$$(m-p)_t = -\underset{(32.38)}{9.09} + \underset{(55.55)}{1.67}\, y_t + \underset{(7.75)}{3.43}\, rp_t - \underset{(6.66)}{1.43}\, ra_t - \underset{(0.81)}{0.09}\, \Delta_4 p_t$$

ADF = -0.26 PP = -1.10	DHS = 46.00 $h_k = 1.11^{***}$
ADF-LS = -1.89 ($\hat{\tau}$=91/III)	PP-LS = -2.28 ($\hat{\tau}$=91/III)
ADF-RS = -6.74^{*} (0.10) ($\hat{\tau}$=89/III)	PP-RS = -6.19 (0.23) ($\hat{\tau}$=89/II)
$\text{ADF-RS}^{\alpha1}$ = -4.78 (0.58) ($\hat{\tau}$=91/II)	$\text{PP-RS}^{\alpha1}$ = -5.24 (0.34) ($\hat{\tau}$=91/II)
$\text{ADF-RS}^{\alpha2,\alpha3}$ = -5.48 ($\hat{\tau}$=89/III)	$\text{PP-RS}^{\alpha2,\alpha3}$ = -5.62 ($\hat{\tau}$=89/II)

Notes: FME:fully modified estimation [Phillips y Hansen (1990)]. t-ratios in brackets. ***, ** and * indicate rejections of the null at 1%, 5% and 10%, respectively. Under certain tests, the levels of significance are reported in brackets.

The ADF, PP and DHS statistics reported in Table 1 are, respectively: the Dickey-Fuller test, the Phillips-Perron test and the Durbin-Hausman test recently proposed by Choi (1994). These are residual-based tests for the null of no cointegration, i.e. tests for the existence of a unit root in the residuals of the regression. Conversely, the h_k test (Leybourne and MacCabe, 1993) directly tests for the null of cointegration[4], i.e the stationarity of the estimated residuals. The

[4] None of these tests has a standard distribution. Critical values for ADF and PP can be found in MacKinnon (1991); the asymptotic critical values for DHS and h_k are tabulated in the references quoted.

reason for including the latter is to provide cover to some degree from the traditional criticism that unit root tests offer scant power in finite samples for rejecting the null. The argument in this connection is that, in a situation in which the economic hypothesis of interest is the alternative one (cointegration), tests will habitually tend not to reject the null (non-cointegration) unless there is substantial evidence against it. Thus, test h_k may be a good supplementary tool to the foregoing ones, whereby if both types of test provide the same inference on the existence or not of cointegration, we may be relatively confident about the robustness of the result.

When the equation is estimated for the whole period, none of the first three tests -ADF, PP and DHS- enables the null hypothesis of non-cointegration to be rejected at standard levels of significance. Likewise, test h_k rejects cointegration at very high confidence levels. This result differs from those obtained in the aforementioned papers and is related, as we will later see, to the inclusion of the most recent years in the sample.

The remaining statistics reported in Table 1 belong to the family of tests proposed in Gregory and Hansen (1993). Starting from the notion of structural change, these authors extend the traditional ADF and PP tests so that they allow, under the alternative hypothesis, the cointegration relationship to change at an unknown point in the sample, either the intercept (LS) or the entire vector of coefficients (RS)[5], including the intercept. In a setting of structural change, in which the power of the traditional ADF or PP tests falls sharply, these tests can detect cointegration while providing an estimation of the date $(\bar{\tau})$ at which the potential break occurs. The following outline -in which the generic name DF has been given to traditional tests and DF^* to those proposed by these authors- may help in the joint interpretation of both types of tests:

[5] The asymptotic critical values of both tests are tabulated in the aforementioned reference.

H_0: non-cointegration	Interpretation	
	cointegratión	**structural change**
DF does not reject DF* does not reject	non-cointegration	-
DF does not reject DF* rejects	cointegration	structural change may be significant
DF rejects DF* rejects	cointegration	no evidence on structural change
DF rejects DF* does not reject	cointegration	no evidence on structural change

Based on the foregoing outline, it is easy to interpret the tests proposed in Gregory and Hansen (1993) as proving useful only when the traditional tests do not enable the null hypothesis of non-cointegration to be rejected. In other cases, when DF rejects, rejection by DF* cannot be interpreted as evidence in favour of structural change, since the alternative contains as a particular case the standard model of cointegration with constant parameters. Likewise, when DF rejects but DF* does not, the result has to be interpreted in the light of the much lesser power of the latter when, in fact, there is no break. In both cases, tests for structural change in regressions with first-order integrated variables as proposed in Hansen (1992) or in Phillips and Quintos (1993) may allow the stability of the cointegration relationship to be evaluated.

One difficulty with the Gregory and Hansen tests lies in the existence, in finite samples, of biases against the null hypothesis (non-cointegration), as evidenced by the results of the Monte Carlo experiments conducted by these authors. To negotiate this obstacle, finite-sample critical values have, where necessary, been simulated[6]. For the RS tests in Table 1 these are: -7.93 (1%), -7.14 (5%) and -6.70 (10%), far higher in terms of absolute values than the asymptotic critical values tabulated in Gregory and Hansen (1993): -6.92, -6.41 and -6.17, respectively.

Only with the ADF-RS test does some marginal evidence in favour of the existence of cointegration appear when the entire vector of parameters is allowed to change. This evidence is even weaker if it is noted that the break point ($\hat{\tau}$) is positioned right at the end of the sample -around the second or third quarter of 1989-, implying that the changes in the long-run parameters are estimated with few degrees of freedom. In any event, the results do appear to suggest that there is a cointegration relationship for the first part of the sample.

This latter observation is substantiated by the evidence given in Table 2, which estimates the same ALP2 demand equation for the period running from the first quarter of 1979 to the last quarter of 1988. The augmented Dickey-Fuller (ADF), Phillips and Perron (PP) and Choi (DHS) tests strongly reject the hypothesis of non-cointegration in the period analysed. Likewise, the Leybourne and MacCabe (h_k) test provides the same inference, not rejecting -at standard confidence levels- the null of cointegration. The remaining tests in Table 2 (F-mean, Lc and F-sup) are those proposed in Hansen (1992), and seek to evaluate explicitly the constancy of the long-run relationship. With none of them is it possible to reject the hypothesis of the existence of a long-run demand for broad money that is stable in the period 1979/I-1988/IV.

Thus, the results to date point to the existence of an ALP2 long-run demand function, which is stable to 1988. However, the

[6] These critical values have been generated using n=5000 replications for T=60 observations. Although n is relatively small for this type of exercise, it is empirically tested that the critical values calculated are sufficiently robust for the purposes of this paper.

Table 2

Estimation method: FME;	**Sample: 1979/I-1988/IV**

$$(m-p)_t = -\ 7.64 + 1.51\ y_t + 3.01\ rp_t - 1.10\ ra_t - 0.41\ \Delta_4 p_t$$
$$\qquad (12.61) \quad (23.03) \quad (9.89) \quad (7.81) \quad (2.87)$$

ADF = -4.89^{**}	PP = -5.20^{**}	DHS = 1119.7^{***}	h_k = 0.061
F-mean = 5.89	Lc = 0.38	F-sup = 16.30 (τ^*= 81/I)	

Notes : See notes to Table 1.

evidence in favour of the cointegration hypothesis when the sample period runs to end-1993 is scant, suggesting a structural break in the ALP2 long-run demand function. It is estimated that this structural change may have occurred around 1989. Moreover, we have also seen how the solution of letting all the parameters change proves unsatisfactory, in the sense that the evidence provided by the RS tests in favour of the hypothesis of two cointegration regimes is weak.

An additional possibility involves restricting the structural change to which the demand for ALP2 would have been subject; that is restricting the type of structural break under the alternative hypothesis in the Gregory and Hansen RS tests to the intercept and some of the slope coefficients. This would improve power properties of the tests and, moreover, it would enable more accurate hypotheses to be analysed on the nature of the possible structural change. Hereafter, several lines of argument used in recent years to signify disturbances that were affecting the demand for ALP are analysed[7].

[7] See, in this connection, the Banco de España 1993 Annual Report or the article Monetary Policy in 1994 in the January 1994 edition of the Banco de España Economic Bulletin.

First of all, since the detection of instabilities in the demand for ALP, it has been occasionally argued that the recent increase in the savings ratio and in resident private-sector wealth may be one of the underlying factors. From this viewpoint, the high long-run income elasticity estimated for this aggregate (1.51 in the equation in Table 2) would capture wealth effects not explicitly taken into account in equation (2). As a result, alterations in the income-wealth ratio would induce instability in the equation estimated and, more specifically, in the parameter that measures long-run income elasticity (α_1).

To test the role played by factors of this type, the ADF-RS$^{\alpha 1}$ and PP-RS$^{\alpha 1}$ tests in Table 1 restrict the possibilities of structural change under the alternative hypothesis to the intercept and to the parameter that affects income in equation (2)[8]. As with the RS tests, the correct critical values were simulated for the sample size at issue, giving the following result: -7.06 (1%), -6.32 (5%) and -6.01 (10%). The results of the tests reported in the table continue not to allow rejection of the null hypothesis of non-cointegration.

Second, it has been suggested that a further factor behind this instability may have been the increase -by historical standards- in the sensitivity of the monetary aggregate to swings in interest rates. In principle, this idea does not challenge the evidence presented that a structural break affecting the demand for ALP occurred in or around 1989. Indeed, at the end of that year, ahead of the Banco de España's far-reaching reform of the reserve requirement ratio in the early months of 1990, the first skirmishes occurred in the so-called "war of the supercuentas" (high-yield deposit accounts). This episode was ultimately framed within a broader process of heightening competition among financial institutions to raise deposits. Similarly, this was the setting of the extraordinary growth of portfolio investment institutions in recent years.

[8] Evidently, this is an indirect test which is conditioned by the absence of a quarterly time series of private-sector wealth for the Spanish economy.

Hence, the ADF-$RS^{\alpha2,\alpha3}$ and PP-$RS^{\alpha2,\alpha3}$ tests reported in Table 1 restrict the structural change to the intercept and to the parameters that affect the own and alternative interest rates in equation (2)[9]. Nor does the evidence drawn from these tests allow the rejection of the null hypothesis of non-cointegration when the alternative allows the change described in the parameters that affect both interest rates in the ALP2 long-run demand equation.

Lastly, a third factor pointed out focuses on the gradual elimination of administrative obstacles to free capital movements and the resulting opening of the Spanish financial system to international markets that considerably broadened the resident private sector's investment and financing opportunities, increasing the range of alternative assets to those included in ALP. From this standpoint, although the main liberalising impulse occurred in 1987[10], 1989 also proved to be an important year, with the lifting of quantitative limits on investment in securities issued on foreign markets[11]. Subsequently, this process steadily firmed through a series of measures that liberalised ecu-denominated deposits (September 1989) and the purchase of securities issued on money markets abroad (April 1990) and allowed residents to hold bank accounts in pesetas and foreign currency at non-resident banks (February 1992). Similarly, the decline in foreign exchange risk derived from the peseta's entry into the ERM on June 19 1989 may also have had a major impact in making foreign markets more accessible to the Spanish financial system.

In accordance with the foregoing, the set of variables in equation (2) has been extended so as to add an interest rate on foreign assets (rx_t). More specifically, the simple average of three-month

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[9] Critical values have not been simulated for this case. However, such values must be greater in absolute terms than those relating to the $RS^{\alpha1}$ tests.

[10] Ministerial Order dated 25-5-87 implementing Royal Decree 2374/1986.

[11] Ministerial Order dated 19-12-88.

Euromarket interest rates on the D-Mark and the dollar have been included in the following estimations[12].

As can be seen in Table 3, this variable is not significant when the equation is estimated for the whole sample 1979/I-1993/IV, and, moreover, the ADF, PP, DHS and h_k tests continue not to allow any amendment to be made to the previous conclusions about the absence of cointegration. On the contrary, the results of the ADF-RS$^{\alpha 5}$ and PP-RS$^{\alpha 5}$ tests do offer some evidence in favour of cointegration, when the possibility of a change in the sensitivity of ALP2 to variations in the foreign rate is envisaged, rejecting[13] at 93% confidence level (the first test) and 85% (the second) the unit root hypothesis in the estimated residuals (non-cointegration). The break point ($\bar{\tau}$) continues to be estimated at around 1989 and, moreover, the foreign interest rate did not prove significant in the previous period.

In keeping with this last observation, Table 4 re-estimates the ALP2 long-run demand equation, including only rx_t values as from the first quarter of 1989 (rx_t^*)[14]. The results in this respect may be qualified as satisfactory, with the foreign interest rate playing a significant role and without there being substantial changes in the remaining parameters -others than the intercept- regarding the estimation to 1998/IV shown in Table 2[15]. The long-run semi-elasticity of ALP2 in relation to variations in the foreign rate is estimated at -1.39, which would imply a semi-elasticity of about -0.7 in relation to individual

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[12] Strictly speaking, these interest rates should be adjusted for exchange rate depreciation expectations. The problem has been obviated with the argument that this component would be I(0), not affecting (asymptotically) the estimation of the long-run parameters.

[13] The critical values simulated for these tests are: -7.30 (1%), -6.65 (5%) and -6.29 (10%).

[14] Using FME here rests on the assumption of rx_t^* being strongly exogenous.

[15] The ADF, PP, DHS and h_k values in this table are for illustrative purposes to enable comparison with the values obtained in the other tables. The relevant tests are ADF-RS$^{\alpha 5}$ and PP-RS$^{\alpha 5}$ in Table 3.

Table 3

Estimation method: FME; **Sample: 1979/I-1993/IV**

$$(m-p)_t = -\underset{(28.25)}{9.08} + \underset{(48.12)}{1.67}\, y_t + \underset{(7.03)}{3.34}\, rp_t - \underset{(5.78)}{1.41}\, ra_t$$

$$- \underset{(0.67)}{0.10}\, \Delta_4 p_t + \underset{(0.009)}{0.001}\, rx_t$$

ADF = -0.25 PP = -1.07 DHS = 40.26 h_k =1.32***

ADF-RS$^{\alpha 5}$ = $\underset{(0.07)}{-6.48^*}$ ($\hat{\tau}$=89/III) PP-RS$^{\alpha 5}$ = $\underset{(0.15)}{-6.07}$ ($\hat{\tau}$=89/II)

Notes: See notes to Table 1.

Table 4

Estimation method: FME; **Sample: 1979/I-1993/IV**

$$(m-p)_t = -\underset{(16.76)}{8.21} + \underset{(29.56)}{1.57}\, y_t + \underset{(9.77)}{2.74}\, rp_t - \underset{(5.97)}{0.85}\, ra_t$$

$$- \underset{(3.00)}{0.34}\, \Delta_4 p_t - \underset{(7.95)}{1.39}\, rx_t^* + \underset{(8.37)}{0.124}\, S89/I$$

ADF=-4.81 PP=-5.54 DHS=564.31 h_k=.0766

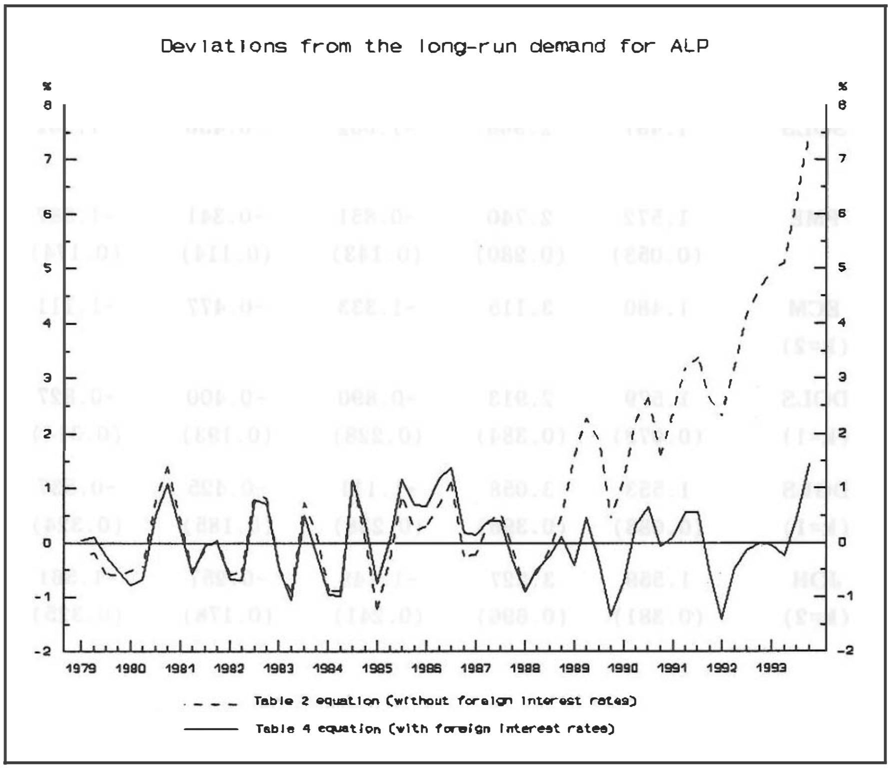
Notes: See notes to Table 1. The variables s89/I and rx_t^* take zero values before 1989 and, thereafter, the former variable takes one, and the latter the simple average of three-month interest rates in Germany and USA.

changes in German and US interest rates.

Nonetheless, the aforementioned results must be viewed with some caution. First, because the foreign interest rate semi-elasticity is estimated with few degrees of freedom (five years' observations) and, therefore, with little accuracy. The robustness of the estimates is analysed below. Second, as the intercept is allowed to vary, the possibility of other unspecific factors taking place from 1989 onwards and leading to a level shift in the demand function cannot be totally dismissed. However, the results from tests LS in Table 1 strongly suggest that such factors are not decisive in getting cointegration.

Chart 1 throws some light on the role played by foreign interest rates, showing the deviations of ALP2 with respect to the long-run paths calculated with the equations estimated in Tables 2 (dotted line) and 4 (solid line). In the period prior to 1989, both lines take very similar paths. As from that year, however, the monetary aggregate clearly tends to grow at higher rates than those compatible with the demand equation estimated with pre-1989 data, and the deviations from the long-run path become increasingly larger and more persistent. Conversely, if the foreign interest rate is included among the determinants of the demand for ALP2 -in other words, if external assets denominated in foreign currency are considered substitutes of those included in ALP- more stationary deviations can be obtained. From this standpoint, the high interest rate differentials in these years would be a basic factor in the high growth of ALP2 in the period. This in turn would be consistent with the small scale of resident portfolio investment abroad and with the increasing weight of the foreign sector in the financing of households and non-financial firms.

Chart 1



Finally, Table 5 shows, along with the FME estimation, the result of the application to the equation in Table 4 of various procedures for the estimation of cointegrating vectors: SOLS (static ordinary least squares: Engle and Granger, 1987), ECM (Hendry and Ericsson, 1991), DOLS (dynamic OLS: Saikkonen, 1991), DGLS (dynamic generalised least squares: Stock and Watson, 1993) and JOH (Johansen, 1988). Of these the latter three, together with FME, are asymptotically efficient according to Saikkonen (1991), i.e. their asymptotic distribution is strictly more concentrated around the true value of the parameter. The details of the estimation are given in the notes to the table.

The estimations of the parameters that affect income, own and

Table 5

Estimation method	y_t	rp_t	ra_t	$\Delta_4 p_t$	rx^*_t
SOLS	1.497	2.968	-1.052	-0.456	-1.382
FME	1.572 (0.053)	2.740 (0.280)	-0.851 (0.143)	-0.341 (0.114)	-1.387 (0.174)
ECM (k=2)	1.480	3.115	-1.333	-0.477	-1.111
DOLS (k=1)	1.579 (0.079)	2.913 (0.384)	-0.890 (0.228)	-0.400 (0.193)	-0.827 (0.314)
DGLS (k=1)	1.553 (0.083)	3.058 (0.398)	-1.111 (0.238)	-0.425 (0.185)	-0.537 (0.324)
JOH (k=2)	1.559 (0.381)	3.227 (0.696)	-1.149 (0.241)	-0.257 (0.178)	-1.561 (0.325)

Note:

Standard errors in brackets (autocorrelation and heteroskedasticity consistent for DOLS). If $x=(y, rp, ra, \pi, rx^*)$ and $z=m-p$, the estimators of the (β) are:

1.OLS: ordinary least squares in the static regression: $z_t = \beta x_t + u_{1t}$

2.ECM: OLS in the regression: $\Delta z_t = \sum_{j=1}^{k} \Delta z_{t-j} \gamma_j + \sum_{j=0}^{k} \Delta x_{t-j} \delta_j - \alpha z_{t-1} + \alpha\beta x_{t-1} + u_{2t}$

3.DOLS: OLS in: $y_t = \beta x_t + \sum_{j=-k}^{k} \Delta x_{t-j} \gamma_j + u_{3t}$

4.DGLS: generalised least squares in regression (3), where it is assumed that u_{3t} follows an AR(1) process.

5.JOH: application of Johansen (1988) to the vector $x=(y, rp, ra, \pi, rx^*)$, in which the variable rx^* is assumed to be weakly exogenous.

alternative interest rates and inflation are fairly stable, irrespective of the estimation method used. On the contrary, the parameter that measures the semi-elasticity of ALP2 to variations in foreign interest rates, though significant in most cases, is that which evidences the greatest variability between estimations, with values ranging from -1.387 with FME to -0.537 with DGLS. That is perhaps a reflection of the limited observations available for its estimation.

Conversely, the estimation of this parameter does prove robust to certain changes in the definition of the relevant foreign interest rate. Thus, for example, if the equation in Table 4 is re-estimated via FME using a weighted average foreign rate[16] instead of the simple average of the rates of the D-Mark and the dollar, the long-run semi-elasticity estimated (-1.279) is very similar to the previous case. Likewise, when German and US rates are included separately in the equation, a semi-elasticity somewhat higher than in the first (-1.0 versus -0.7) is obtained, although some accuracy is lost in the estimation.

3. CONCLUSIONS

This paper set out to ask whether it was necessary, and in which sense, to modify the conclusions other authors had formulated on the Spanish demand equation for broad money (ALP), mainly those relating to the existence of cointegration and to the stability of the coefficients defining the long-run relationship. The empirical evidence analysed in the paper in this connection suggests that the shifts witnessed in recent years in the demand for this monetary aggregate have, far from being transitory episodes confined to the short term, substantially altered its long-run properties. Thus, it is not possible to accept, at reasonable confidence levels, the existence of an ALP demand relationship such as that estimated in Cabrero et al. (1992) which is stable in the period 1979:I-1993:IV. This counters the results obtained for the period

[16] Specifically, the weighted average rate used is that appearing in the first column of Table 19.24 of the Banco de España's "Boletin Estadístico".

immediately prior to 1989 and, along with the more specific tests reported here, suggests the existence of a structural break around 1989.

Regarding the direction of the structural change, analysis has been made of the empirical relevance of some of the explanations which, in various short-term economic reports, have been used to signal disturbances that were affecting the course of the monetary aggregate: the increase in the private-sector savings ratio; the heightened sensitivity of ALP to changes in the own and alternative interest rates; and the growing openness of the Spanish financial system to international markets as the obstacles to free capital movements were progressively lifted. It is the third of these three hypotheses which appears to have the most prominent role when explaining the apparent instability of the traditional ALP demand equations. Thus, only when the returns other than those included in ALP are widened to incorporate, as from 1989, foreign interest rates is it possible to recover a long-run demand equation for ALP with any appearance of stability in recent years. However, a word of caution: the estimated value of the parameter that measures the semi-elasticity of ALP demand to foreign interest rates evidences, due to a problem of degrees of freedom, great variability between different estimations. Nonetheless, irrespective of the estimation procedure used, it does prove possible to reject at high levels of confidence that this parameter is not significant.

As noted earlier, the paper focuses on the parameters defining the ALP long-run demand equation. Nonetheless, this should not obscure the fact that the description of the process of dynamic adjustment to the long-run path may very often be of prime interest -when a forecasting tool is required, for instance. In a different context, factors not considered here, especially expectations about an exchange rate depreciation, can play an important role in helping interpret short-run ALP developments. This was the case, for example, in 1992 when in the quarters prior to the successive devaluations of the peseta there was a marked, albeit transitory, delocalisation of deposits.

In general, the broadness of this monetary aggregate -which incorporates highly heterogeneous financial assets in terms of the liquidity services they provide- makes its behaviour relatively difficult to interpret, since ALP demand combines strictly transactional motives with speculative motives.

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