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Second Thoughts on Second Moments

Panel Evidence on Asset-Based Models of Currency Crises

The evidence is broadly supportive of an asset view of speculative attacks and the importance of the variance of monetary aggregates in predicting currency crises, but it cast some doubt on existing theories.

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Summary findings

The literature on speculative attacks has been given new impetus by the collapse of the European currency arrangements beginning in 1992, by the Mexican peso crisis and after-effects in 1994, and most recently by speculative attacks across Asia.

One strand of this literature stresses the importance of imbalances in stocks of monetary and financial aggregates rather than traditional "flow" factors, arguing that massive, volatile capital flows have become a dominant feature of the global landscape, and that exchange-rate levels and current accounts have not proved convincing as proximate causes of crises.

Galindo and Maloney test two popular asset-based models of speculative attacks — Krugman and Rotemberg (1992) and Calvo and Mendoza (1995) — especially their emphasis on the second moments of monetary aggregates.

Analyzing monthly panels of appropriate countries in three regions, they find evidence for the importance of money/reserve ratios predicted by both models, and their variance as predicted by Calvo and Mendoza.

But the variance of velocity does not appear to be important, casting some doubt on the Krugman-Rotemberg target zone framework and the interpretation of the Calvo-Mendoza results.

This paper — a product of the Poverty and Economic Management Unit of the Latin America and the Caribbean Region— is part of a larger effort in the region to understand the determinants of macroeconomic instability. Copies of the paper are available free from the World Bank, 1818 H Street NW, Washington, DC 20433. Please contact Marta Cervantes, room I8-095, telephone 202-473-7794, fax 202-522-0054, Internet address mcervantes@worldbank.org. William Maloney may be contacted at wmaloney@worldbank.org. June 1998. (27 pages)

Second Thoughts on Second Moments

Panel Evidence on Asset-Based Models of Currency Crises

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I. Introduction

The literature on speculative attacks has been given new impetus by the collapse of the European currency arrangements beginning in 1992, the Mexican Peso crisis and after effects in 1994, and most recently by attacks across Asia. A comprehensive review of the numerous approaches and findings to date is offered by Kaminsky, Lizondo, and Reinhart (1997).

One strand of this literature stresses the importance of imbalances in stocks of monetary and financial aggregates rather than traditional "flow" factors, arguing that massive and volatile capital flows have become a dominant feature of the global landscape, and exchange rate levels and current accounts have not proved convincing as proximate causes of crises (see, for example, Calvo and Mendoza, 1995). The earliest genre of these model dates from Salant and Henderson (1978) and Krugman (1979) and have been further elaborated by Flood and Garber (1984) and Obstfeld (1986) among others. In most, a persistent and monetized budget deficit leads to an offsetting fall in reserves. Forward looking investors, anticipating the eventual abandonment of the peg, attack the currency when the remaining stock of reserves equals the decline in domestic money demanded that will occur when the currency floats. This provides a rationale for the inclusion of ratio of reserves to money (Bilson 1979, Edwards 1989, Kaminsky and Reinhart 1996, Klein and Marion 1994, Sachs, Tornell and Velasco 1996) rather than the more traditional scaling by imports (see for example, Edin and Vredi, 1993, Frankel and Rose, 1996) or GDP (Collins 1995) and the inclusion of the rate of growth of domestic credit (Edwards, 1989, Frankel and Rose, 1996).

However, as Calvo and Mendoza note, in the Mexican case, the apparent fiscal surplus in 1993 appears to contradict this type of speculative attack model. They argue that the focus should rather be on the stochastic evolution of demand for monetary aggregates, particularly M2. Demand

for domestic assets by foreign capital or for private expenditure can suddenly evaporate leaving the monetary authorities the choice of using sterilized intervention and weakening the currency, or risking the collapse of a weak banking system. Noting that the log of the ratio of M2 to reserves appears to follow a random walk, Calvo argues that its higher volatility in Mexico raises the probability of wandering into crisis above what it would be in Austria with a comparable reserve ratio.

However, the Calvo-Mendoza view shares a closer kinship with the literature on target zones beginning with Williamson (1985), Frenkel and Goldstein (1986), and Krugman (1991), than with the domestic credit driven models that they critique. Krugman and Rotemberg (1992) developed the theoretical bridge to the speculative attack literature and derive the specific conditions under which a band cannot be defended and a crisis may be expected.[1] The target zone framework is more appropriate to both the European case and the major Latin American countries in the 1990's which were more often than not managing their exchange rates within a band. It is also rich in predictions about the role of reserve to money ratios and the second moments of monetary variables in generating crises.

Despite the popularity of both these views, to date there has been no systematic testing of their predictions and particularly about the importance of the volatility of monetary aggregates. This paper attempts to do so and finds only partial support. Beginning with the target zone literature, we generate a set of testable specifications and show their broad similarity to the Calvo Mendoza view. We compile a data set focused on a sample of countries in the late 1980s and 1990s whose exchange rate arrangements and capital account regulations are appropriate to the model. We then employ

[1] Edin and Vredin (1993) and Ötker and Pazarbaşioğlu (1994)also analyze attacks on target zones but with in a very different framework that does not yield predictions about second moments.

panel estimators that preserve the temporal dimension often lost in previous studies that pool observations. We then test the predictions of both the target zone and the Calvo-Mendoza framework.

We tightly restrict the number of variables included in the regressions. This is primarily because our goal is to test the importance of a few heretofore unexamined variables, rather than to predict crises per se. However, it is also the case that the choice to work over a relatively short interval at high frequency necessarily implies that many of the variables that have appeared important in other studies are unavailable for many or all the countries in the sample at the required frequency.

I. Theoretical Background from the Target Zone Literature

In the Krugman-Rotemberg framework, the nominal exchange rate is assumed to follow a simple, although standard, log linear monetary model of the exchange rate

$$s = m + v + \frac{\eta E[ds]}{dt} \qquad (1)$$

where s is the log of the spot exchange rate, m the log of domestic money supply, v a shift term capturing shocks to money demand including those to real income, velocity etc., and the expected rate of depreciation times η, the interest semi-elasticity of money demand. The term v is assumed to evolve as a random walk with drift:

$$dv = \mu dt + \sigma dz \qquad (2)$$

where z is a wiener process: $dz \sim N(0,1)$, and μ is the rate of drift. Money supply is assumed to be

passive and altered only to keep the exchange rate within the target zone. As the exchange rate moves toward the end of the band, intervention by the central bank reduces the money supply to maintain s in bounds and the smooth pasting equilibrium holds. However, as in the earlier literature, an attack will occur if the stock of reserves is eroded to where it equals the decline in the demand for money that would result from the collapse. Krugman and Rotemberg show this quantity

$$m' - m = -1/\lambda \tag{3}$$

where

$$\lambda = \frac{-\eta\mu + \sqrt{\eta^2\mu^2 + 2\eta\sigma^2}}{\eta\sigma^2} \tag{4}$$

This implies that an attack occurs when

$$\frac{R}{D+R} < 1 - e^{-1/\lambda} = \tau \tag{5}$$

The threshold ratio of reserves to high powered money below which an attack occurs, τ, is a function of η, μ, and σ^2. An increase in the drift or the variance of the shocks to money demand, or an increase in the sensitivity of money demand, expected depreciation, through the interest rate lowers the threshold. The difference between the first and second elements above can be seen as an index of proximity to the threshold that holds the promise of being a useful predictor of attacks.

Scaling the threshold by the money multiplier, equation (5) is broadly consistent with Calvo's focus on the log of M2 over reserves converted into domestic currency and its variance, rather than

that of velocity. The two measures of variance diverge to the degree that purchasing power parity fails and reserves are not proportional to income. In the estimations that follow, we first test explicitly the Krugman-Rotemberg specification, and then the Calvo hypothesis.

III. Estimation

We construct panels of up to nine years of monthly data for 14 countries. We choose this frequency first because it seems appropriate given the rapidity with which fundamentals can change. Second, it generates enough degrees of freedom to permit focusing on a restricted period, 1987-1995, which corresponds reasonably well to the assumptions of the model: high degrees of short term capital flows, reasonably open economies, and authorities committed to maintaining a target zone or, in the limit, a peg as determined by the IMF publication, *Exchange Arrangements and Exchange Restrictions*. The downside of this approach is that the availability of indicators at this frequency sharply restricts the range of countries that can be included, and the span of data available. This is especially the case for Asia where only Korea and Malaysia publish industrial production numbers, the only feasible proxy for the output variable required to calculate velocity and to estimate interest elasticities. Since the latter managed a target zone for only a brief period, we exclude the Asian region from this part of the work and do not include the1997 crises. In total, our sample includes nine European countries -Austria, Denmark, France, Italy, Holland, Finland, Ireland, Portugal and Spain- and five Latin American countries -Brazil, Chile, Colombia, Mexico, Argentina -for which equation (5) can be tested directly. Since the Calvo-Mendoza hypothesis does not require output measures to calculate relevant variables, in the second section we can employ a much broader range of countries however, to remain comparable with the first section, again, we do not address

the most recent crises.

We generate an index of speculative pressure similar to that of Eichengreen and Wyplosz (1995). Reserve movements and real exchange rate are standardized by their standard deviations and combined. As in Sachs, Tornell, and Velasco, and Kaminsky, Lizondo and Reinhart, interest rates were not included in the index due to sharp movements in Latin America that are often unrelated to attacks.

The results we present are those leaving the index as a continuous measure. The literature frequently discusses the incidence of "speculative pressure" that often falls short of a full blown attack (see Svensson, 1994). Such episodes, though falling below whatever arbitrary cut off is employed to define a discrete "crisis" are arguably driven by similar dynamics. It may therefore be inefficient to discard this information, create a dichotomous variable, and then employ limited dependent variable techniques to infer the determinants of the underlying continuous latent variable. On the other hand all measures that weight innovations by the country-specific standard deviation treat a one standard deviation of the index as equally important episodes of speculative pressure, whether in Austria or Mexico. This is defendable to the degree that countries differ in "normal" movements in reserves or exchange rates, and hence also in what should be considered a crisis. But as an alternative, we also create a binary crisis index informed by movements in the index, but modified by what the literature recognizes as legitimate speculative attacks. All the analysis was run using this variable both with the complete sample, and truncating each country series immediately after a crisis to focus on the run-up to each event. However, perhaps due to the limited number of crises relative to observations, no specification appeared remotely significant and we do not report the results.

The series for the variance and trend in velocity movements are generated in two ways. First, individual GARCH models were fit for each of the 14 countries. The trend was derived as the forecast of a time series model for the log of velocity of general form:[2]

$$\Delta\log(v_t)=\alpha_0+\alpha_1\log(v_{t-1})+\sum_{i=1}^{n}\alpha_{i+1}\Delta\log(v_{t-i})+\epsilon_t \tag{6}$$

As conventional, we assumed that the error term is normally distributed with mean zero and variance h_t, where

$$h_t=\gamma_0+\sum_{i=1}^{p}\gamma_i\epsilon^2_{t-i}+\sum_{i=1}^{q}\beta_i h_{t-i} \tag{7}$$

For every country, an acceptable individual specifications was generated that removed residual GARCH effects, usually with a GARCH(1,1) specification or a ARCH(1). Second, to provide a smoother alternative under the assumption that longer term volatility may enter speculators' decisions, we construct a six month rolling variance of $\Delta \log(v)$ and use its six month moving average for the trend.

To generate a consistent set of interest semi-elasticities, a simple model of M1 in first differences was estimated using two stage least squares. In all cases, the coefficient on money was of the correct sign, and almost always significant. Since the available interest rates are implicitly those paid on assets often included in M2, the estimated semi-elasticities using this

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[2] For a more general discussion of GARCH models see Bollerslev, Engle and Nelson (1993). The inclusion of a "levels" effect in the mean equation is now common in estimations of continuous time stochastic volatility models. See, for example, Andersen and Lund (1997) for an application to short term interest rates.

aggregate were, unsurprisingly, very often positive. In the absence of data on returns on less liquid assets, this means the specification can only be run with M1.[3] Since we are concerned only with the direct effect of depreciation on money demand through the interest rates, cointegration based estimation methods were not appropriate since they generate the total impact elasticity through all variables in the system.[4] Although the literature on estimating interest response of money demand is long and contentious, as we will see, the precision of these estimates does not appear critical to the results. The threshold value is scaled by the money multiplier, so as to make it consistent with the ratio of reserves to M1, rather than base money.

Finally, unlike the European subsample, the Latin American countries adopted a target zone or peg at different times within our sample period. Further, in the case of Argentina, the deceleration of inflation in the early part of the stabilization plan introduced a high degree of both real exchange variance, interest rates and other variables that was unrelated to the sustainability of the peg per se. We therefore begin the sample in 1992:1 when inflation was falling to levels below 50%. The effect in both cases is to generate an unbalanced panel.

As a preliminary test of the model, table 1a presents the thresholds calculated first using the GARCH and then the moving average specifications of the variance, as well as the level of R/M1 for the entire sample, and table 1b the same information for several countries experiencing crises in both Europe and Latin America. As is immediately evident, on average, the reserve ratio

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[3] This also raises questions about the interpretation of the interest rate coefficient in empirical tests of monetarist models of the exchange rate that employ M2, and use these same interest rates as the opportunity cost of holding it. Estimates available on request.

[4] Johansen(1995) and Lütkepol (1994) argue that the coefficient from cointegrating regressions cannot be interpreted as the necessary partial elasticities since they capture shocks transmitted through all other variables and cannot be allowed a *ceteris paribus* interpretation.

is far above the threshold and that, even at their maxima, these thresholds are very low. In the months before crises, only in the case of Italy was the reserve to M1 ratio remotely close to either threshold. In general, a strict interpretation of equation (5) would imply thresholds that tend to be so low that we should virtually never see a crisis.

One possible conclusion might be that this arises from the inaccuracy of our estimates of the elements of equation(5). However, figure 1 shows the value of the threshold to be very insensitive to even large movements around our estimates.[5] First, since velocity series are either I(1) or I(0), differencing them leaves them stationary and, not unexpectedly, with essentially no drift, μ. Very large increases would be needed to raise τ to .1 even at values of σ^2 an order of magnitude greater than the maxima observed.[6] Given the relatively standard tools employed, it seems unlikely that our estimates are off by these magnitudes. At current levels of σ^2 and μ, even large differences in the interest semi-elasticity have very little effect. Again, since our estimates are of similar orders of magnitude to those found elsewhere, this is unlikely to be the problem. Given reasonable values for the arguments involved, the literal application of this model is unlikely to generate crises at the reserve ratios generally observed.

This, of course, is not in itself evidence against the target zone framework more generally for analyzing speculative attacks. The Krugman-Rotemberg model is admittedly heuristic in intent and departs from a simple monetary model of the exchange rate that has persistently resisted empirical verification. Nonetheless, it is not unreasonable to expect that the arguments

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[5] The average multiplier for the sample is used to scale the threshold.

[6] In an earlier application to Colombia, Mexico and Germany, Carasquilla (1995) found much higher thresholds. This was due, however, to unusually high estimates of the drift term.

in eq (5) appear in some form among the determinants of currency crises. Our estimation strategy is therefore first to take the model literally, and then progressively to loosen the constraints on the underlying arguments until the final regression is

$$Pressure_t = \beta_0 + \sum_{i=1}^{n} [\beta_{R/M,i} \frac{R}{M1}_{t-i} + \beta_{\mu,i}\mu_{t-i} + \beta_{\sigma,i}\sigma_{t-i}] + \beta_\eta \eta. \quad (8)$$

Table 2 presents the results of these regressions. Columns 1a and 1b present the specification with only the proximity to the threshold index calculated using the GARCH estimates of the variance and drift, and the moving average estimates respectively. Columns 2a and 2b allow R/M1 and τ to enter separately, again calculating the latter using the two separate measures of variance and drift. Columns 3a and b estimate (8) above, unconstraining the arguments in τ.

The results offer only partial support to the model. Standard Hausman and Breusch - Pagan tests dictate using either pooled or variable effects estimators, depending on the subsample. In each case, an equal number of lags for all variables were included and the lag structure was pared down to where the last set of lags was insignificant. Contemporaneous values were excluded since in a crisis situation, we would expect a large shock to reserves would be reflected in R/M1. In virtually all cases, only two lagged sets of variables were significant. The sum of the coefficients are reported and the probability value of the F-tests on their joint significance below.

Virtually all specifications show F or χ^2 tests on the overall significance of the regression significant below the 8% level and for Europe and the overall sample, below the 5% level. In all cases, the proximity to the threshold index enters with the anticipated sign, and significantly,

regardless of the variance and drift measures employed. Of concern, however, is that when the index is broken into the asset ratio and the threshold, τ, the latter enters with the predicted sign in the European sample, but is significant only for the GARCH specification at the 10% level. The reserve ratio, on the other hand, emerges of the predicted sign and very high levels of significance in virtually all specifications. This suggests that to the degree that the index was significant, it was driven largely by the reserve ratio.

Disaggregating τ into its component parts, the drift term, μ, enters with correct sign and significantly at the 10% level in the GARCH specifications for the European and the complete samples, but insignificantly or of the wrong sign for all other specifications. The σ terms , are also of the anticipated sign in roughly half the specifications and enter at the 11-13% level only in the European specifications, as with drift, with the correct sign. The semi-elasticity of money demand also shows unstable signs and never enters significantly. In sum, the only specifications for which the variance and drift terms enter consistently with the model and of some significance are the European specifications. However, these are also the only specifications for which the asset ratios enter with the wrong sign. When the semi-elasticity is dropped from the regression, the sign reverses to that anticipated although both drift and variance terms become slightly with the latter now significant at the 15% level only. The other regressions largely unaffected (results available on request).

The highest level of explanatory power, as measured by the R^2 is for Latin America, at only 7.2% of the variance explained. Further, to remove the possibility that the estimates of interest elasticities were driving the aggregated specifications, they were also run with a common value of .1. However, consistent with the discussion above, this had essentially no impact on the

results.

IV. Test of the Calvo-Mendoza View with an Expanded Sample

Calvo and Mendoza argue that ln(R/M2) and its variance should appear as important in determining speculative attacks. Since we no longer calculate velocity or estimate semi-elasticities, we do not need measures of economic activity and the sample can be expanded to include countries previous dropped for lack of data. The sample now includes four Asian countries- Indonesia, Korea, Malaysia, and Thailand. We also group in this category, "Asia+," Israel which, while clearly not sufficient as a category of its own, is an important case study for target zones.[7] To the five existing Latin American countries we add Uruguay and to Europe we add Greece and the UK and Sweden for the M2 regressions. The African countries in the Franc zone were not included despite their long-standing peg to the French currency since capital flows remain largely restricted. We employ the moving average representation of the variance rather than estimate 24 individual GARCH specifications.

Table 3 presents the ratio of reserves to M2, its log, and the standard deviations of the latter across the sample period employed in the regressions. Figure 2 presents the evolution of these variables across a longer period for a selection of countries. What is immediately clear from both is that geographical generalizations are not robust. As Calvo points out, Mexico does have a much higher variance of R/M2 relative to Austria, and this may offset the fact that it has a higher reserve to M2 ratio. But the other two Latin countries hit in 1994-95, Argentina and

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[7]Williamson (1996) has a detailed analysis of the crawling bands of Chile, Colombia and Israel.

Brazil have roughly the same degree of volatility as, and significantly higher reserve ratios than Austria, as well as every Asian country with the exception of Malaysia. At the time of the Tequila crisis, Colombia and Chile had levels of variance similar to those of Austria. Overall, volatility in Latin America would be difficult to distinguish from Europe and reserve ratios are, on average, higher.

It is true that, in table 3, the moderate Latin American volatility arises partially from having dropped the high inflation periods in Argentina, Brazil and Mexico. We defend this on two grounds. First, it can be argued that these are unusual periods and thus do not share the same data generation process as the other countries in the panel. Second, the variances across these periods dwarf the *relatively* small rises around the tequila period and in preliminary regressions tended to generate the inverse correlation with crises from that predicted. These high variances may be "real" but they may possibly arise if large increases in money supply, and the expected proportional depreciation of the currency are not coincident.

Figure 2 suggests some support for the Calvo-Mendoza hypothesis. Chile, Colombia and Uruguay, countries largely unaffected in the Tequila episode, had extremely low variances across this period while Brazil and Mexico, with relatively high reserve ratios, showed rises in their variances in the early part of 1994 to among the highest levels in the sample. On the other hand, Italy and Spain at the end of the 1994 showed comparable levels of variance but with much lower reserve ratios and yet experienced no crisis while Argentina showed low variance and relatively high reserves and was still hit.

Table 4 presents the results of regressing the pressure index on the log of reserve ratios and their variance. For comparison with the previous section, we begin working with M1. F-tests

suggest 6 lags of the two variables. As before, the asset ratio is significant for the entire sample, and Europe and for Latin America at the 7% level. However, the variance is now significant for the whole sample, Europe, and Latin America although it enters with incorrect sign in the latter and in Asia+. The variables taken together are statistically significant for all except the Asia+ regression, although again, the overall explanatory power is under 5% of the variance.

The results improve if we work with R/M2 as suggested by Calvo and Mendoza. The sample size increases for Europe because Sweden publishes M2 and the U.K. publishes a proxy for M2 (the retail component of M4) but neither publish M1. R/M2 is of the predicted sign for all but Asia+ although it is now not significant within Latin America. The variance is very significant and of the correct sign for all except Asia+. Again, all the regressions, with the exception of Asia+ are very significant.

The poor performance of the model for Asia+ may results from two factors. First, since in none of the countries was there a true speculative attack across the sample period, the movements in the standardized index may represent noise unrelated to speculative pressure. Moderate depreciations designed to preserve competitiveness in Korea, or Israel will get very large weight, yet occur in relatively healthy macro-environments. The fact that the model predicts so poorly in this case may be considered support for it overall. It also suggests that, for the other regions, the index is not just picking up noise. It may also be, however, that despite the loosening of capital controls over time, some countries, like Korea, still managed short term flows and therefore do not correspond well to the model.

The fact that the variance now enters with the correct sign in the Latin subsample is supportive of the variance of M2/R being the more appropriate of the two monetary aggregates.

The explanatory power also increases in every case except Asia+. This raises the question of whether the relative success of the Calvo-Mendoza model compared to the Krugman-Rotemberg model is solely due to using M2 rather than M1. As empirical studies of monetary models of the exchange rate frequently employ M2, this might have been a more desirable aggregate to employ in section III were it not for the unavailability of corresponding interest elasticities. As an alternate test, in table 5 we present the results of a specification analogous to that of Calvo-Mendoza, where the variance of R/M2 is replaced by the variance of the inverse of velocity, PY/M2. As in the more complete regressions using M1, the results are not supportive of the Krugman-Rotemberg specification: the variance of the velocity does not enter significantly in any regression and the signs are the opposite of those predicted in both the overall and European regressions.

This finding provokes some second thoughts about the more successful Calvo-Mendoza approach as well. The shocks to broad money demand that it postulates as critical to bringing on crises should presumably also show up in the variance of velocity yielding similar empirical findings. The fact that they do not raises the question of what is driving the significance of the variance of R/M2, the variance of M2, or of reserves. This is not necessarily bad news. Finding that the second moment of reserves helps predict crises is still useful information for policy makers even if not entirely in line with the formal motivation in terms of shocks to M2. A possible concern is that if in the run up to a crisis, reserve losses become progressively larger, this may show up both in the pressure indicator, that has as one component the change in reserves, as well as in lags of the variance of M2/R. Attempting to eliminate this problem by running a probit with the binary crisis index capturing recognized attacks, as in section III,

yielded insignificant results. However, as before, this may be due to the few crises relative to observations.

Conclusions:

The paper provides some evidence in favor of an asset view of speculative attacks and the importance of the second moments of monetary aggregates in predicting crises. In the regressions for both the Krugman- Rotemberg target zone model and the Calvo-Mendoza approach, the stock of money relative to reserves appears very significant and of the predicted sign in most specifications. The results for the drift and variance terms for the innovations in velocity are less consistently supportive of the first model with only the GARCH specifications for Europe and the overall sample generating the predicted signs and borderline significance. These results cannot be seen as strong evidence in favor of the target zone framework or as offering much confidence in the elusive measure of proximity to crisis that it theoretically offers. The variance of reserves to the money aggregates suggested by the Calvo-Mendoza approach, however interpreted, appears more significantly and may contribute additional explanatory power to models seeking to predict crises.

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Appendix I: Data

M0: All from line 14 of IFS statistics except: Italy 1995 from Banca D'Italia, Economic Bulletin, Number 22, Feb 1996.Colombia, Banco de la República.

M1: line 34 IFS or, if unavailable, M1. Colombia, Banco de la República. Not available for U.K. or Sweden.

M2: line 34 + line 35 quasi money, IFS. UK: Building Societies pay interest on demand deposits. In 1989, some became banks. Bank of England continued calculating M1 without new banks

until 1990. Then stopped. M2 is the retail component of M4, The Bank of England suggested the omitting very large depositors who had some market power and who were unlikely to be using M4 for transactions purposes was the correct measure. 1987-1994 from Bank of England, Statistical Abstract, 1995 part 2 Detailed monetary statistics. 1995 values provided by the Bank of England. Colombia, Banco de la Republica.

Money multipliers: Ratios of M1 to Base money. Chile, Boletin Mensual, Banco Central de Chile.

Industrial Production: All from line 66 of IFC except: Portugal: June 1994- Boletím Mensal de Estadistica, Instituto Nacional de Estatistica, Bank of Portugal; Chile, Boletín Mensual, Banco Central de Chile; Colombia, Banco de la República, Brazil, Banco Garantía; Argentina;

Real Exchange Rate: IFS line reu or rec. For Latin America, EP*/P where P* is weighted average of the WPI of the principal commercial partners of each country. US WPI.

Interest rates: IFS line 60l or closest market determined rate. Argentina, Informe Económico, Ministerio de Economía y Obras y Servicios Públicos.

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Appendix II: Countries and Sample Periods

Asia: Israel 1987:1-95:12,Indonesia 1987:1-95:12, Korea 1988:12-1995:12; Malaysia 1987:1-1988:12.

Europe: Austria, Denmark, France, Italy, Holland, Finland, Greece, Ireland, Portugal, Spain, Sweden, UK; all 1987:1-1995:12

Latin America: Argentina 1992:01-95:12; Brazil 1994:07-95:12; Chile 1987:1-95:12; Colombia 1987:1-95:12; Mexico 1991:11-95:12; Uruguay 1990:12-95:12.

Table 1a: Complete Sample Summary Statistics

	Mean	S.D.	Max	Min
Pressure Index	-0.000	0.025	0.444	-0.319
τ_1	0.016	0.022	0.212	0.000
τ_2	0.019	0.021	0.169	0.001
R/M1	0.706	0.710	3.562	0.054
σ^2_1	0.003	0.005	0.042	0.000
μ_1	-0.004	0.056	0.369	-0.249
σ^2_2	0.007	0.015	0.180	0.000
μ_2	-0.002	0.023	0.070	-0.190
η	0.684	0.688	2.010	0.069

Notes: τ_1 = threshold, σ_1 =variance of innovations, μ_1= drift using GARCH. τ_2, σ_2, μ_2 using moving average. R/M1 = reserves in domestic Currency divided by narrow money. η = semi-elasticity of money demand.

Table 1b: Thresholds and Reserve Ratios for Selected Speculative Attacks

Country	Crisis -1	τ_1	τ_2	R/M1
Argentina	1994:11	0.036	0.036	0.885
Brazil	1994:11	0.072	0.024	1.658
Mexico	1994:11	0.010	0.007	0.322
Finland	1992:08	0.009	0.001	0.165
Italy	1992:08	0.031	0.012	0.054
Spain	1992:08	0.023	0.014	0.382
Portugal	1992:08	0.011	0.010	1.146
Spain	1992:10	0.034	0.015	0.356
Portugal	1992:10	0.014	0.017	0.947
Spain	1993:04	0.025	0.031	0.299
Portugal	1993:04	0.018	0.015	0.828
Spain	1993:06	0.004	0.032	0.323
Portugal	1993:06	0.010	0.015	0.826
France	1993:06	0.001	0.002	0.119

Notes: τ_1 = threshold using GARCH, τ_2 using moving average. Crisis-1= month before crisis.

Table 2: Determinants of Speculative Attacks on Target Zones, 1987-1995

	ALL						EUROPE						LATIN AMERICA					
	1a	1b	2a	2b	3a	3b	1a	1b	2a	2b	3a	3b	1a	1b	2a	2b	3a	3b
Index$_1$ E-03	-4.37						-0.302						-7.82					
	0.000						0.014						0.000					
Index$_2$ E-03		-4.28						-0.243						-7.87				
		0.000						0.047						0.000				
R/M 1 E-03			-4.26	-2.70	-4.56	-4.23			-0.788	-0.674	0.35	0.375			-7.44	-5.38	-7.31	-3.75
			0.000	0.000	0.000	0.000			0.035	0.039	0.041	0.028			0.000	0.010	0.000	0.007
τ_1 E–02			-0.613						8.15						-6.4			
			0.366						0.100						0.664			
τ_2 E–02				-8.61						4.48						-14.8		
				0.082						0.697						0.287		
μ_1 E–02					1.41						3.85						-0.63	
					0.070						0.066						0.218	
μ_2 E–02						-3.56						0.73						-1.88
						0.010						0.438						0.337
σ_1 E–01					1.21						2.09						-5.22	
					0.638						0.109						0.840	
σ_2 E–01						-0.117						0.863						-5.75
						0.667						0.128						0.267
η E-04					-1.2	-2.96					5.89	4.7					6.33	-7.840
					0.915	0.785					0.420	0.5					0.144	0.874
constant *E-03*	2.78	2.770	2.87	3.37	2.71	2.97	0.721	0.691	0.076	0.322	-0.405	-0.326	9.310	9.310	11.2	11.90	10.90	10.50
	0.007	0.007	0.008	0.002	0.066	0.027	0.498	0.512	0.951	0.813	0.723	0.762	0.088	0.090	0.068	0.040	0.144	0.126
R^2	0.045	0.041	0.043	0.046	0.044	0.047	0.010	0.010	0.012	0.010	0.017	0.017	0.054	0.050	0.048	0.054	0.072	0.059
Obs	1091	1091	1091	1091	1091	1091	828	828	828	828	828	828	263	263	263	263	263	263
Overall Signif.	26.59	24.65	13.29	14.08	8.16	8.67	8.47	6.13	11.04	7.24	14.14	14.79	8.40	7.96	4.32	4.77	2.83	3.33
P value	0.000	0.000	0.000	0.000	0.000	0.000	0.015	0.047	0.026	0.124	0.048	0.039	0.000	0.000	0.002	0.001	0.007	0.002

Notes: Results are the summation of the estimated parameters of the first two lags of the variables and beneath them, the P-value of tests under the null hypothesis that the coefficients are jointly equal to zero. Specification "a" uses the GARCH estimates of the variance and drift while specification "b" employs a six month moving average. Tau is the threshold level of R/M1 at which an attack would be expected. The sample size is 1987:1- 1995:12 for Europe while those for Latin America depend on the individual country (See appendix II). The complete sample and Latin American models are pooled regressions while the European specification employs a random Effects estimator as dictated by Hausman and Breusch-Pagan tests. F tests for the pooled regressions and Chi squared tests for the random effects estimators were used to evaluate the significance ofthe estimated parameters and of the overall regression.

Table 3: Sample Means and Standard Deviations

Country	R/M2	Log(R/M2)	S.D. Log(R/M2)
Europe			
Austria	0.08	-2.57	0.16
Denmark	0.12	-2.12	0.23
Finland	0.12	-2.14	0.22
France	0.06	-2.83	0.16
Greece	0.15	-2.04	0.51
Ireland	0.28	-1.28	0.21
Italy	0.07	-2.67	0.31
Netherlands	0.09	-2.43	0.21
Portugal	0.26	-1.42	0.36
Spain	0.13	-2.06	0.22
Sweden	0.17	-1.82	0.39
United Kingdom	0.07	-2.62	0.17
Latin America			
Argentina	0.27	-1.31	0.14
Brazil	0.24	-1.44	0.18
Chile	0.50	-0.74	0.27
Colombia	0.45	-0.84	0.25
Mexico	0.20	-1.67	0.32
Uruguay	0.14	-2.05	0.33
Other			
Israel	0.16	-1.82	0.19
Indonesia	0.18	-1.73	0.19
Korea	0.16	-1.85	0.14
Malaysia	0.30	-1.20	0.09
Thailand	0.22	-1.52	0.23

Notes: moments correspond to sample period used in estimations (See appendix 2).

Table 4: Tests of the Calvo-Mendoza Hypothesis,1987-1995

Sample:	All	Europe	Latin America	Asia+
M1				
Log(R/M1) E-03	-2.03	-0.439	-17.0	-2.04
	0.004	0.039	0.070	0.478
V[Log(R/M1)] E-02	4.77	11.2	-3.76	-10.8
	0.003	0.000	0.001	0.646
Constant E-03	-0.258	-1.05	-6.41	-0.649
	0.001	0.142	0.106	0.446
R^2	0.018	0.049	0.080	0.000
Observations	1577	960	232	289
Overall Significance	3.39	49.06	2.67	0.69
	0.000	0.000	0.002	0.763
M2				
Log(R/M2) E-03	-2.47	-0.764	-3.16	0.044
	0.001	0.037	0.150	0.894
V[Log(R/M2)] E-02	7.43	15.4	13.7	-1.82
	0.000	0.000	0.000	0.591
Constant E-03	-5.99	-2.19	-12.1	4.29
	0.004	0.177	0.220	0.530
R^2	0.029	0.061	0.111	0.000
Observations	1769	1152	232	289
Overall Significance	5.38	74.34	3.41	0.55
	0.000	0.000	0.000	0.879

Notes: We report the summation of the first six lags of each variable and below it the P-Value of the test under the null that the coefficients are jointly equal to zero. All models are estimated as pooled regressions except the European ones where Hausman and Breusch-Pagan tests dictated a random effect model. F-tests for the pooled and Chi squared for the random effects regressions were used to evaluate the overall significance of the specification.

Table 5: Test of Simplified Krugman Model, 1987-1995
(Calvo-Mendoza with variance of velocity)

Sample:	All	Europe	Latin America
	M2		
Log(R/M2) E-03	-3.35	-0.732	-6.75
	0.000	0.001	0.044
V[(YP)/M2] E-02	-2.38	-2.89	31.8
	0.689	0.188	0.837
Constant E-03	-6.99	-0.622	-1.54
	0.003	0.736	0.207
R^2	0.028	0.030	0.056
Observations	1192	960	232
Overall Significance	3.24	2.72	1.34
	0.000	0.001	0.198

Notes: We report the summation of the first six lags of each variable and below it the P-value of the test under the null that the coefficients are jointly equal to zero. All models estimated are estimated as pooled regressions. F tests were used to evaluate the overall significance of the specification.

Figure 1: Sensitivity of τ to variance, drift, and semi-elasticity of money demand

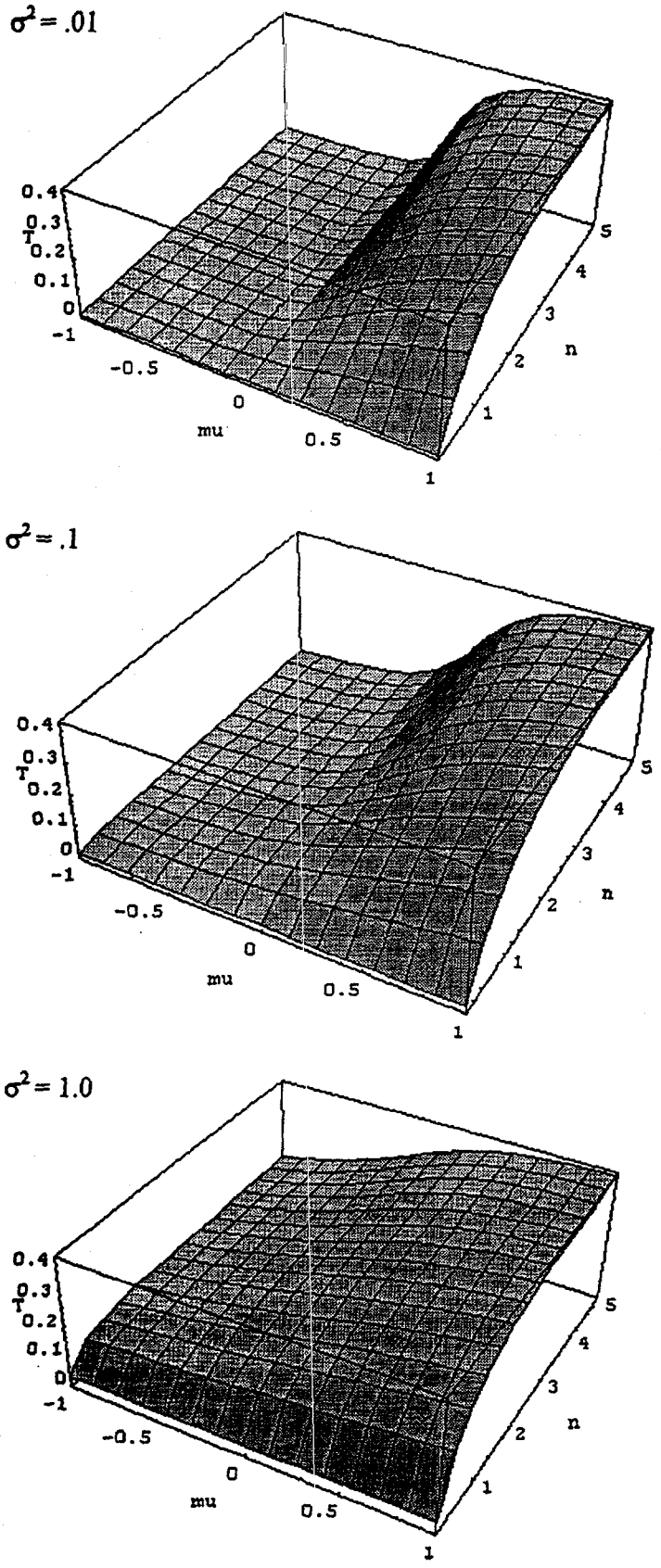


Figure 2: R/M2 and VAR(ln(R/M2)

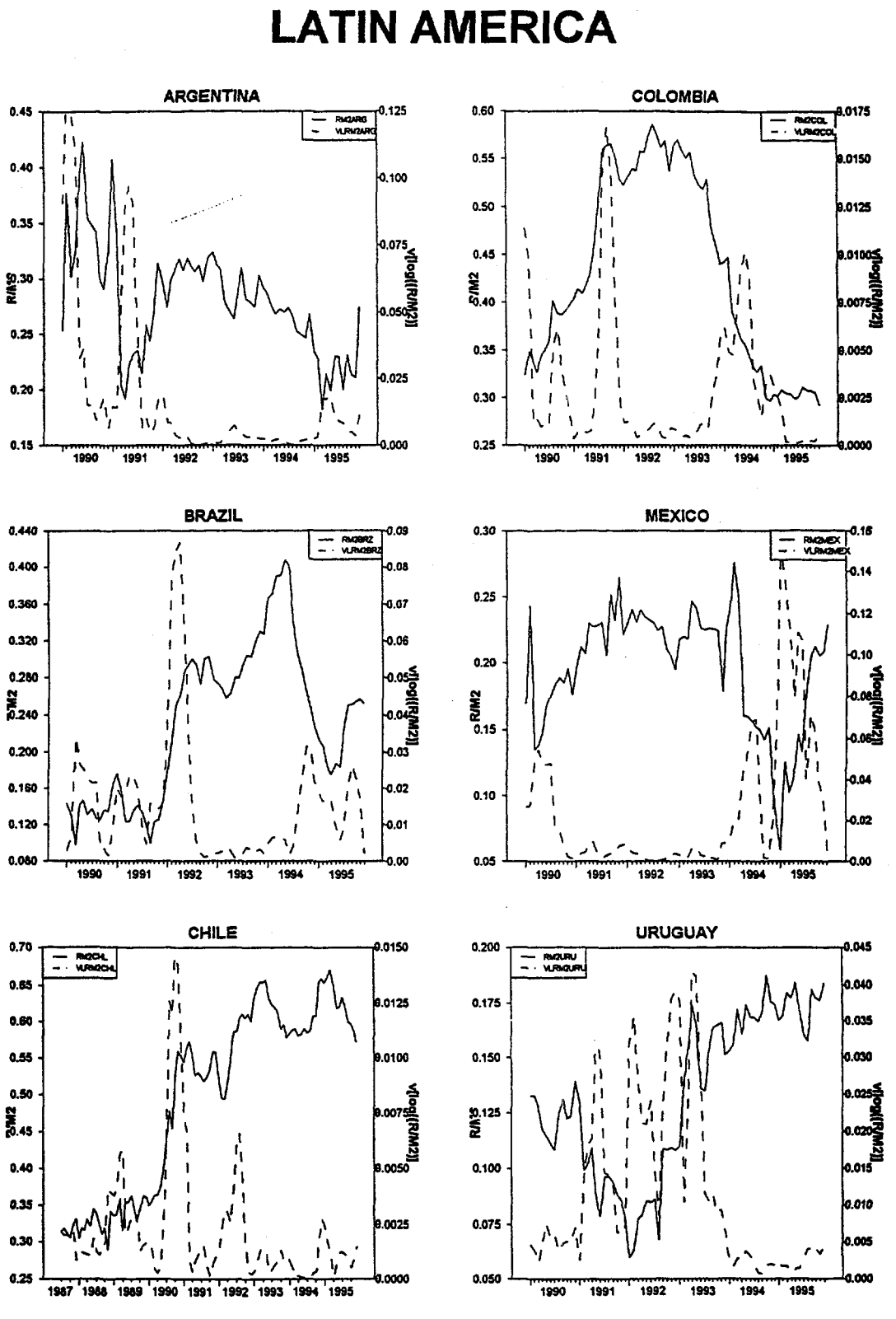
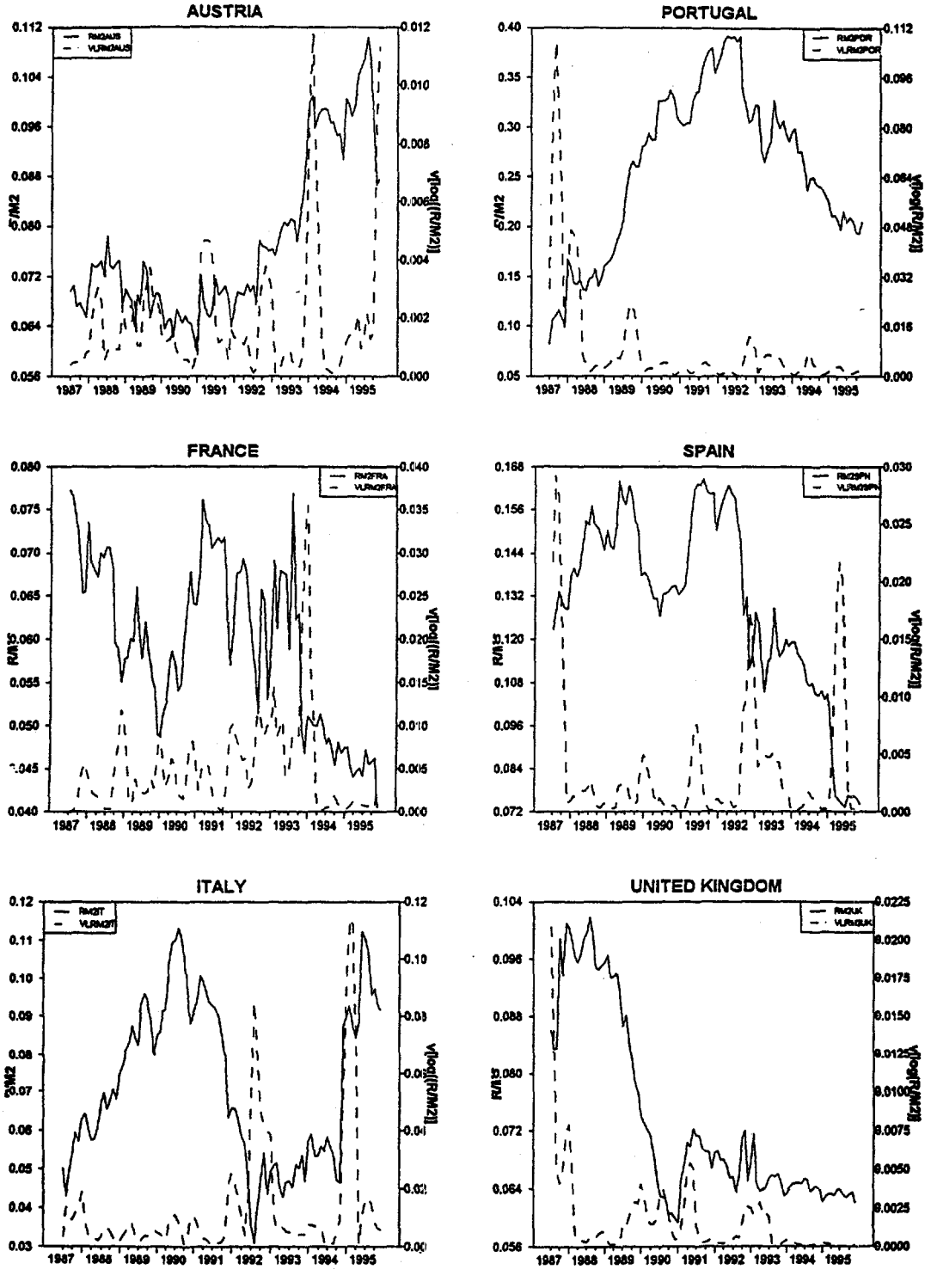


Figure 2, cont.

EUROPE



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