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Macroeconomic Fundamentals and the DM/$ Exchange Rate: Temporal Instability and the Monetary Model

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Editorial

On April 19 - 20, 2001 the Oesterreichische Nationalbank sponsored a Workshop organized by Richard Clarida (Columbia University), Helmut Frisch (TU Wien) and Eduard Hochreiter (OeNB) on „Exchange Rate and Monetary Policy Issues". It took place at the Institute for Advanced Studies, Vienna. A number of papers presented at this workshop is being made available to a broader audience in the Working Paper series of the Bank. This volume contains the forth of these papers. The first ones were issued as OeNB Working Papers No. 44, 46 and 47.

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Macroeconomic Fundamentals and the DM/$ Exchange Rate: Temporal Instability and the Monetary Model

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1 Introduction

Despite the volumes written on the failure of empirical exchange rate models, there is a key feature of the record that is much overlooked, namely there are some subperiods of floating rates during which macroeconomic models provide reasonable explanations of monthly or quarterly exchange rate movements and other subperiods during which the explanatory power of these models completely disappears. Frankel [1979], for example, found an overshooting model of the DM/$ exchange rate fitted remarkably well in-sample during the mid-1970s. But when Frankel's sample was updated to include the late 1970s and 1980s, researchers found a lack of cointegrating relationships (e.g., Meese and Rogoff [1988] and Boothe and Glassman [1987]) and parameter estimates that were either insignificant or significant and of the wrong sign (e.g., Frankel [1983,1984] and Backus [1984]). By the 1990s, more powerful dynamic approaches to estimating and testing for cointegrating relationships had been developed (e.g., Johansen [1988]) and Phillips [1991]), but again some studies found evidence of cointegrating relationships (e.g., MacDonald and Taylor [1994], MacDonald and Marsh [1997] and Cushman, Lee and Thorgeirson [1996]), while others using different sample periods did not (Baillie and Pecchenino [1991] and Papell [1987]). This temporal inconsistency is also prevalent in out-of-sample forecasting exercises. The seminal study of Meese and Rogoff [1983] showed an inability of exchange rate models to outperform the random walk model in out-of-sample forecasting over a sample that included the early 1980s. When this study was updated to include the 1990s and the use of dynamic models, some researchers found the structural models outperformed the random walk model at shorter forecasting horizons (e.g., MacDonald and Taylor [1994] and MacDonald and Marsh [1997], while others using similar methodologies but different sample periods found the random walk still dominated (e.g., Chinn and Meese [1995].[1]

The temporal inconsistency of exchange rate models to explain floating rates provides compelling support for the view that one empirical exchange rate model with fixed coefficients is very unlikely to perform well either in sample or out of sample. Meese remarks that "the most menacing empirical regularity that confronts exchange rate modelers is the failure of the current generation of empirical exchange rate models to provide stable results across subperiods of the modern floating rate period (Meese [1986], p.365)." And yet this empirical regularity is largely ignored by studies documenting the failure of empirical exchange rate models.

The purpose of this paper is to reexamine whether macroeconomic fundamentals matter for the short-run movements of the DM/$ exchange rate in the context of the monetary model. We first explore the nature of the empirical

[1] Meese observes that "A few empirical researchers have estimated structural models that can predict currency movements better than the random walk over particular subsamples of the modern floating rate period. But on the whole, forecasting success with empirical exchange rate models has proved to be ephemeral (Meese [1990], p. 126).

regularity that exchange rate models experience temporal instability. Using recursive techniques we find that a composite monetary model of the DM/$ exchange rate experiences structural breaks on seven occasions over a sample that begins in March 1973 and runs through December 1998.[2] Given the observed break points, we are able to isolate four subperiods (or regimes) during which the null of no structural change cannot be rejected. These exchange rate regimes involve the following subperiods: 1) July 1974 through August 1978 (regime 1); 2) October 1979 through August 1984 (regime 2); 3) September 1987 through December 1993 (regime 3); and 4) January 1994 through November 1998, the end of the sample (regime 4).[3]

Second, we investigate the presence of cointegrating relationships among the variables of the monetary model. To this end, we use the cointegrating VAR framework of Johansen [1988] and estimate unrestricted VAR models over the entire floating rate period as well as over the four subperiods characterized by "relative" parameter constancy. The results for the full sample are consistent with the results of the structural change analysis, in that the properties of the errors of the full-sample statistical model are found to be inconsistent with the properties of the underlying statistical theory. Errors appear to be non-normal, ARCH, heteroscedastic, and serially correlated despite increasing the order of the system. However, once the unrestricted model is re-estimated for each of the four subperiods involving relative parameter constancy, the properties of the errors markedly improve in two out of four subperiods (regimes 1 and 4). In the remaining two subperiods (regimes 2 and 3) errors continue to exhibit non-Gaussian behavior suggesting that the monetary model provides an inappropriate statistical model in which to test structural hypotheses in these subperiods.

Consequently, we focus our cointegration analysis on the two subperiods in which the unrestricted statistical model does not seem to be incongruent with the underlying statistical theory. For each of these subperiods we find multiple cointegrating vectors, at least one of which enters the exchange rate equation. Thus, although the statistical properties of the monetary model are problematic when based on the full sample and two of the four subperiods of floating, our cointegration results indicate macroeconomic fundamentals matter for the movements of the DM/$ exchange rate in at least some subperiods of the floating rate experience.[4]

[2] This section of the paper builds on and extends the analysis in Goldberg [1991] and Goldberg and Frydman [1996a].

[3] Other studies that find the monetary model to be temporally unstable include Boughton [1987], Meese and Rogoff [1988] and Goldberg and Frydman [1996a,b]). Also, the studies of Engel and Hamilton [1990] and Kaminsky [1993], among others, indicate that the process governing exchange rate movements undergoes discrete switches. Although these studies provide univariate analyses, their results are suggestive that the multivariate process governing exchange rates exhibits switching behavior.

[4] One explanation of the inability to fit unrestricted models over the full sample and in regimes 2 and 3 is the dominance of non-fundamental factors over fundamental factors in

Although the finding of cointegration suggests support for the monetary model with RE, further analysis of the cointegrating relations reveals a very striking result: different sets of macroeconomic fundamentals are significant during different subperiods of floating, i.e., even when macroeconomic fundamentals matter for short-run exchange rate movements, the way they matter changes considerably from one time period to the next.[5] We argue that this finding is inconsistent with the monetary model with RE.[6]

In a forthcoming paper we offer an explanation of our finding that different sets of macroeconomic fundamentals matter during different time periods (Frydman and Goldberg [2002]).[7] This explanation makes use of an alternative approach to modeling of expectations — the theories consistent expectations (TCE) framework.[8] The TCE framework incorporates the most important insight of RE, that agents use theories to look forward in forecasting exchange rates. But it relaxes the most controversial aspect of RE, that agents forecast as if they knew the precise quantitative magnitudes of the parameters of the one true model. Instead, TCE assumes that agents are endowed with imperfect knowledge of how the exchange rate is related to macroeconomic fundamentals. TCE is based on the fact that there exists a pluralism of theories describing exchange rate dynamics and that these theories can at best provide agents with qualitative knowledge about the variables and parameter signs that are potentially relevant in forming short-run exchange rate expectations.

The remainder of the paper is structured as follows. Section 2 explores the nature of the temporal instability of the exchange rate process within the context of the monetary model. The section finds that the exchange rate process is episodically unstable, involving different sets of significant macroeconomic fundamentals during different subperiods of floating. Section 3 argues that the monetary model with RE cannot be reconciled with these structural change findings. Section 4 provides concluding remarks and discusses how the monetary model with TCE can be reconciled with the finding that different sets of fundamental variables matter during different time periods. .

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driving short-run movements of exchange rates during these periods. See Frydman and Goldberg [2001a] for a discussion of the relevance of non-fundamental factors for exchange rate movements.

[5] This finding provides empirical support for an earlier observation by Meese that "A perusal of the published empirical work reveals that the set of explanatory variables most correlated with exchange rate movements depends on the sample period analyzed (Meese [1990], p. 126)." For earlier empirical evidence on this point see Goldberg [1991] and Goldberg and Frydman [1996a].

[6] It is tempting to rely on a Lucas critique explanation, but we show in Frydman and Goldberg [2001b] that the Lucas critique does not allow for the variables entering the equilibrium exchange rate equation to change from one time period to the next within the context of the monetary model.

[7] Also see the working paper version of this paper (Goldberg and Frydman [2001a]).

[8] The TCE framework and its rationale were developed and extensively discussed in our earlier work (Goldberg [1991], Goldberg and Frydman [1993,1996b] and Frydman and Goldberg [2002]). The original idea of theories consistent expectations was proposed in Frydman and Phelps [1990].

2 The Temporal Instability of the Monetary Model

Although the exchange rate literature provides compelling evidence of the temporal instability of monetary models (see footnote 3), the nature of this instability remains unclear. There are a number of possibilities: 1) the same set of macroeconomic fundamentals matter for exchange rates in every time period, but the influence of these variables as measured by regression coefficients changes over time; 2) different sets of macroeconomic fundamentals matter for exchange rates during different time periods; and 3) there are time periods during which no macroeconomic fundamentals appear to be significant. The purpose of this section is to explore which of these possibilities is consistent with the data over the modern floating rate period.

2.1 Testing for Structural Change

We begin by testing the following composite monetary model of the exchange rate for structural instability over a sample that begins in March 1973 and runs through June 1999 for the DM/$:

$$s = \beta_0 + \beta_1 m_r + \beta_2 y + \beta_3 y^* + \beta_4 i + \beta_5 i^* + \beta_6 \pi + \beta_7 \pi^* + \beta_8 k_r \qquad (1)$$

where s denotes the mark price of dollars, m_r denotes domestic minus foreign money supply levels, y is the domestic income level, i is the domestic short-term interest rate, π is the domestic expected secular inflation rate, k_r denotes domestic minus foreign cumulative trade balances and a "*" denotes foreign value. (See the data appendix for a description of the data.) Equation (1) is well known and encompasses the flexible-price model of Frenkel [1976] and Bilson [1978,1979], and the sticky-price models of Dornbusch [1976], Frankel [1979] and Hooper and Morton [1982]. The only atypical feature of the exchange rate model in equation (1) is the absence of most of the usual symmetry restrictions. We maintain the symmetry restrictions on the money supply and cumulative trade balance variables in order to conserve degrees of freedom when conducting our in-sample regression analysis in the next two subsections.[9]

In order to test the exchange rate model in equation (1) for temporal instability, we make use of two single-equation procedures, the CUSUM test and the Quandt ratio (QR) technique.[10] The cusum test is used to establish in a statistical sense that a break has occurred and the QR technique is used to determine

[9] It should be noted that with RE, either nominal interest rates or secular inflation rates enter the cointegrating exchange rate equation, but not both. We show in Goldberg and Frydman [2001a] that the reduced form of the monetary model with TCE does give rise to a cointergrating relationship for the exchange rate in which both nominal interest rates and secular inflation rates enter, and where the interest rate variables enter without the symmetry restriction. On the absence of the symmetry here see also Goldberg [2000].

[10] The structural change analysis follows the approach in Goldberg [1991,2001] and Goldberg and Frydman [1996a,b]. See Brown, Durbin and Evans [1975] for a discussion of the CUSUM and Quandt ratio techniques. Boughton [1987] also uses the cusum test in testing for structural change.

the most likely location of the break point. Note that these procedures test recursively for the possibility of one or more break points, rather than relying on tests that choose break points a priori. They therefore provide a mechanical, non-discretionary way for locating points of parameter instability.[11] There are several reasons for using the CUSUM test. First, it is based on OLS residuals and so is easily implemented. Second, Ploberger and Krämer [1996] show that the CUSUM test is valid in the presence of general trends in the regressors. Finally, this single-equation approach allows us to focus on structural change occurring in the cointegrating exchange rate relation.[12]

It is important to point out that structural change is not an easy matter to test for or to model, especially when the objective is to locate subperiods within which parameter estimates are relatively constant. Any finding of specific periods of stability will depend not only on the particular structural change tests employed in the analysis, but also on the assumed size of the type I and II errors. In order to check the robustness of our results we varied the type I and II errors in our analysis. As expected, this affected the particular subsamples of relative parameter stability. However, it is important to emphasize that the nature of the structural change findings is robust to such changes in the test procedure; it remains true that different sets of macroeconomic fundamentals are significant during different subperiods of floating.

The results of our structural change analysis are reported in figure 1, where break points are indicated by solid vertical lines. Figure 1 shows that the monetary model experiences structural breaks on seven occasions over the sample period, giving rise to four sufficiently long subperiods of relative parameter constancy. The first subperiod or exchange rate regime begins in July 1976 and runs through September 1978; the second regime begins in November 1979 and runs through August 1984; the third regime begins in September 1987 and runs through December 1993; and the fourth regime begins in January 1994 and runs until the end of the sample in January 1998. Again, this evidence of episodic structural change is consistent with the findings of earlier studies (see footnote 3).[13]

[11] It should be noted that we do not use the Quandt ratio technique as a test for structural change, but rather use this technique to locate the most likely point of the break. Thus, the problems discussed in Hansen [2000] of applying the Quandt ratio technique as a test of parameter instability in the presence of general trends in regressors does not apply.

[12] There are systems-based test procedures, such as Hansen and Johansen [1993]. But these procedures are parameter intensive, and severely limit the ability to test for episodic structural change.

[13] It is interesting to note that the sample period used in Frankel [1979] begins in July 1976 and ends in February 1978. Our structural change findings provide justification for Frankel's decision to omit the period immediately following the collapse of fixed rates from his sample.

2.2 Estimating Statistical Models

In order to test for cointegration using the approach of Johansen [1988] and Johansen and Juselius [1990,1992], statistical models with well behaved (Gaussian) errors need to be estimated in the form unrestricted VAR's. We therefore estimate unrestricted VAR's for the full sample and for the four exchange rate regimes and check whether these models have reasonable statistical properties, in particular whether the errors are Gaussian.

The Johansen [1988] framework involves estimating the following vector error-correction model (ECM):

$$\triangle Z_t = \Pi Z_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \triangle Z_{t-i} + \Phi D_t + \epsilon_t \tag{2}$$

where $Z_t = \{\mathrm{s}\ \mathrm{m}_r\ \ \mathrm{y}\ \ \mathrm{y}^*\ \ \mathrm{i}\ \ \mathrm{i}^*\ \ \pi\ \ \pi^*\ \ \mathrm{k}_r\}$ is a column vector of observations on the current values of our n=9 variables in the model, $\triangle$ denotes first-difference operator, Π is a n$\times$n matrix of equilibrium parameters[14], Γ is a n$\times$n matrix of short-run-adjustment parameters, D_t is a vector of deterministic terms and ϵ_t is a vector of random Gaussian errors. The rank of the matrix Π is of central importance because according to the Granger representation theorem, if the rank of Π (r(Π)) is such that 0<r(Π)<n-1, then there exists r(Π) cointegrating vectors and the matrix Π can be expressed as a product of two matrices, $\Pi=\alpha\beta'$, where α is a n$\times$r matrix of equilibrium adjustment parameters and β is a n$\times$r matrix of r(Π) cointegrating vectors.

In estimating unrestricted VAR's, two decisions need to be made: 1) the nature of the deterministic variables (intercepts, trends and seasonals) and whether intercepts and/or trend coefficients should be restricted; and 2) the order of the VAR. In terms of deterministic components, Doornik, Hendry and Nielsen [1998] show that including an unrestricted constant and a restricted trend in the analysis is a good idea because erroneously omitting such deterministic variables leads to substantial mis-specification bias, whereas including these variables works well in terms of good power and reasonable size even when the true data generating process (DGP) possesses no trend and no drift. Doornik, Hendry and Nielsen [1998] show that allowing for an unrestricted trend is correct if one exists, but leads to substantial size distortion when the underlying DGP possesses a restricted trend. Pesaran and Smith [1998] support these conclusions and find that the restricted trend model is preferable when one or more of the underlying variables are trended, which with income levels and money supplies

[14] We use the label "equilibrium" instead of the more common lablel "long-run" in describing the parameters of the Π matrix. This is because with TCE, the equilibrium component of the model in ΠZ_{t-1}provides only a temporary anchor towards which the short-run movements of the exchange rate move (given by ΔZ_t). Given the temporary nature of this equilibrium anchor, which is due to the shifting nature of expectations functions, we refer to this anchor as a medium-run anchor. We discuss this aspect of TCE in Goldberg and Frydman [2001b] and Frydman and Goldberg [2002].

seems likely. In light of these findings, we allow for an unrestricted constant and a restricted trend (case 2* in Osterwald-Lenum [1992]).

As for choosing the order of the VAR, a common practice in the literature is to start with some "high "order and then use model selection criteria such as Akaike Information Criterion (AIC), Schwarz Bayesian Criterion (SBC) or an F-statistic to test down to the most congruent specification. The objective is to eliminate serial correlation with a minimum number of lags, since each lag quickly eats up degrees of freedom.[15] Unfortunately the Johansen [1988] framework is parameter intensive (each lag requires nine additional parameters for the monetary model estimated here), so that the strategy of estimating separate models for each exchange rate regime severely limits the number of lags that can be tested.

One of the important findings of this study, however, is that departures from the Gaussian framework, in terms of serial correlation, lack of normality, heteroscedasticity and ARCH errors are all less severe once the existing structural change is taken into account. In fact, a reasonable statistical model could not be found when one model with fixed coefficients was estimated over the entire sample period, even when many lags were added to the specification. In contrast, reasonable statistical models were found for two of the four separate exchange rate regimes.

Given the degrees-of-freedom limitation in the separate exchange rate regimes, the strategy of choosing the order of the systems began with an AR(1). If no or minimal departures from the Gaussian framework were detected at normal significance levels, then an AR(1) was chosen. This was the case for exchange rate regimes 1 and 4. As for exchange rate regimes 2 and 3, departures from Gaussian errors with an AR(1) specification were problematic. We tested higher orders up to an AR(4) without success. As for the full-sample statistical model, working up to an AR(8) did not deliver a congruent model.

Table 1 presents these findings for the full-sample VAR model as well as the VAR models estimated over the four subperiods of relative parameter constancy. The columns labeled C refer to VAR's estimated for the full composite model of nine variables, whereas the columns labeled P for exchange rate regimes 1 and 4 refer to VAR's estimated for parsimonious models that excluded all insignificant variables. We discuss the parsimonious models below. Four tests are reported. ARCH tests are based on the F-form and tests for heteroscedasticity are based on White [1980]. Tests for normality are based on the Doornik and Hansen [1994] chi-square test, which has been shown to be suitable for small samples. Tests for serial correlation are based on an F-statistic with one-to-four lags for exchange rate regimes 1, 2 and 4 and with one-to-five lags for exchange rate regime 3. Significance levels are shown in parentheses, where A, H, N and

[15] Cheung and Lai [1993] find that serial correlation is a serious problem for the Johansen approach and that the usual lag selection criteria such as AIC and SBC may be inadequate.

S denote significant levels of ARCH, heteroscedastic, non-normal and serially correlated errors.[16] Non-normal errors may not be a serious problem, since Johansen's method requires only i.i.d. errors.[17] But heteroscedastic, ARCH and serially correlated errors are a problem (e.g., see Eitrheim [1992] and Cheung and Lai [1994])

Column 2 in table 1 reveals that the statistical model for the full sample is not well behaved. Of the nine equations of the system there are: 1) four equations with significant ARCH errors at the .05 level or above; 2) two equations with significant heteroscedastic errors at .10 and .05 levels; 3) eight equations with significant departures from normality at the .01 level; and 4) four equations with significant serial correlation, three at the .01 level and one at the .10 level.

Columns 3-8 in table 1 present the results for the separate subperiods of relative parameter constancy. The estimated models for exchange rate regimes 2 and 3 are still problematic. Although allowing for structural change eliminated all significant ARCH errors and most of the departures from normality in both regimes, serially correlated errors still remain in four equations of the model for exchange rate regime 2 at the .01 level and in six of the equations of the model for exchange rate 3, two each at the .01, .05 and .10 levels.[18]

In terms of the models for exchange rate regimes 1 and 4, the errors from the full composite models are much improved. For the model of exchange rate regime 1, there are two equations with ARCH errors at the .10 level, no evidence of either heteroscedasticity or departures from normality and two equations with significant serial correlation, one at the .05 level and one at the .10 level. For the model of exchange rate regime 4, there is no evidence of ARCH errors, one case of heteroscedasticity at the .01 level, two cases of significant non-normal errors at the .01 and .10 levels and one case of significant serial correlation at the .10 level.

Given the departures from the Gaussian framework for the statistical models of the full sample and exchange rate regimes 2 and 3, we do not test any structural hypotheses in these sample periods. Although not perfect, the errors of the statistical models for exchange rate regimes 1 and 4 are reasonable. We estimated higher order VAR's for these exchange rate regimes in an attempt to eliminate all significant departures from Gaussian behavior, but this failed

[16] In estimating the statistical model for exchange rate regime 3, an unrestricted dummy for the period of 1990:06 through 1991:01 was used in order to account for the structural change in the German money supply variable. It should be mentioned that departures from the Gaussian framework for the statistical model of exchange rate regime 3 remained problematic without the money dummy.

[17] Gonzalo [1994] shows that maximum likelihood estimators of the cointegrating vectors are robust to non-normal errors, although Cheung and Lai [1993] find that the trace test is more robust to non-normality than the max eigenvalue test.

[18] Tests for heteroscedasticity on the AR(3) models for these two exchange rate regimes could not be run due to a lack of degrees of freedom.

to improve on matters. Given the limitation on the degrees of freedom in the separate exchange rate regimes, we decided to use the AR(1) specification in testing structural hypotheses for exchange rate regimes 1 and 4. Hence, taken as a whole, the diagnostic results reported in table 1 indicate that allowing for structural change, although not leading to perfectly clean errors, can improve matters substantially.

2.3 Cointegration and the The Nature of the Temporal Instability

With reasonable statistical models in exchange rate regimes 1 and 4, we are able to test structural hypotheses on the equilibrium components of the models. We focus on two questions in this subsection: 1) Do macroeconomic fundamentals matter for exchange rate movements?; and 2) Do different sets of macroeconomic fundamentals matter for exchange rates during different subperiods? The first question can be addressed by testing for the number of cointegrating vectors (i.e. testing the rank of Π) and checking whether any of these cointegrating vectors affect the short-run dynamics of the exchange rate (by testing weak exogeneity restrictions on the first (exchange rate) row of the α sub-matrix of Π).

It is important to address the issue of whether cointegration analysis makes sense for subperiods of four to five years of monthly data. First, from a statistical standpoint we have no choice. The exchange rate process undergoes episodic structural change and it makes no sense to presume an unchanging process and estimate a model with fixed coefficients over the entire sample. This is borne out not only in the structural change findings of section 2.1 but also in the poorly behaved errors of the full sample VAR of section 2.2.

Second, cointegration tests are well known for their low power against near cointegration alternatives (i.e., against processes with slow rates of mean reversion), which has motivated a number of researchers to employ very long data sets (e.g., see Abuaf and Jorion [1990] and Lothian and Taylor [1994]). But this lack of power would be a problem for us only if we were unable to reject the null of no cointegration. The fact that we are able to reject the null with only four to five years of monthly data actually strengthens our results and suggests that the rate of mean reversion in exchange rates and prices is not so slow.

Last, but not least, it is not unreasonable to assume that exchange rates and goods prices mean revert fairly quickly to their equilibrium anchors, certainly within four or five years. This assumption appears to be inconsistent with a large body of evidence suggesting that if exchange rates and goods prices mean revert to PPP, they do so very slowly (with half lives of between four and five years).[19] We show in Goldberg and Frydman [1993,1996b] and Frydman and

[19] See Froot and Rogoff [1995] for a survey article on this issue.

Goldberg [2002], however, that with TCE, the cointegrating anchor towards which exchange rates and goods prices mean revert may or may not imply a movement towards PPP. We use this finding in Goldberg and Frydman [2001b] to show that half-life calculations from PPP regressions should not be interpreted as capturing the speed with which exchange rates and goods prices mean revert.

The trace (TR) and max eigenvalue (MAX) test statistics of Johansen [1988] and Johansen and Juselius [1990] for the rank of Π are well known. Table 2 reports the findings of these tests. Again, we report the results for the full composite models (columns C) and the parsimonious models (columns P). All test statistics are adjusted for degrees of freedom by multiplying the test statistics by (T-nm)/T, where T denotes sample size, n the number of endogenous variables and m the order of the system. This small-sample correction is important because the Johansen test statistics are biased in small samples toward finding cointegration too often if asymptotic critical values are used (see Reimers [1992] and Cheung and Lai [1993]). If the .10 significance level is used as the rejection criterion, then both the TR and MAX tests indicate the presence of at least one cointegrating relationship in each of the two exchange rate regimes.

As is sometimes the case, the MAX and TR statistics lead to different conclusions concerning the number of cointegrating vectors in the two exchange rate regimes. Focusing on the full composite models and the .10 cutoff, the table shows one and three cointegrating vectors based on the MAX test and three and four cointegrating vectors based on the TR test in exchange rate regimes 1 and 4 respectively. Although we discuss the parsimonious models below, we note here that for these models, the MAX and TR tests continue to indicate the presence of cointegrating relationships, with two and four vectors based on the MAX test and three and four vectors based on the TR test for exchange rate regimes 1 and 4 respectively. These results are consistent with Johansen and Juselius [1992] and Pesaran and Smith [1998], among others, which also find the MAX test to identify fewer cointergrating vectors then the TR test.

Testing additional structural hypotheses on the cointegrating spaces of our models is conditional on imposing some rank for Π, and we report below results based on the rank results of the TR test. The rationale is that the power of the MAX (and the TR) test is known to have low power against near cointegration alternatives (e.g. see Eitrheim [1992]), so that the results of the TR test are not inconsistent with those of the MAX test, i.e., a rejection is much more informative than a non-rejection. Also, Cheung and Lai [1993] find that the TR test is more robust to nonnormality than the MAX test and there are some departures from normality in the two exchange rate regimes. It is important to note, however, that the results of testing for weak exogeneity and exclusion restrictions on the β matrix lead to the same conclusion irrespective of whether we rely on the rank results of the TR or MAX tests: different sets of macroeconomic variables are significant during different time periods.

Although the results in table 2 indicate the presence of multiple cointegrating relationships, they do not indicate whether any of these cointegrating relationships enter the exchange rate equation of the system. This question can be addressed by testing whether the exchange rate is weakly exogenous, i.e., by testing zero restrictions on the first row of the equilibrium adjustment matrix α (see Johansen and Juselius [1990]). These results are given in table 3. For the first exchange rate regime, the assumption of weak exogeneity can be rejected easily for both the full composite model and the parsimonious model. In exchange rate regime 4, weak exogeneity is rejected for the full composite model at the .07 level. When the insignificant variables are dropped from this model, weak exogeneity can be rejected at near the .01 level. The results in table 3 show that at least one of the cointegrating relationships found in each exchange rate regime enters the exchange rate equations of the models, suggesting there are cointegrating relationships that are relevant for the short-run movements of the exchange rate.

Although fundamentals seem to matter for short-run exchange rate movements in exchange rate regimes 1 and 4, this does not necessarily imply that they matter in a way suggested by the monetary model with RE. One way to address this issue is to examine whether different sets of macroeconomic variables are significant in regimes 1 and 4. As we argue in Frydman and Goldberg [2001a,b] and below, the finding that different sets of fundamentals matter for exchange rates during different regimes is inconsistent with the monetary model with RE.

This question can be addressed by testing whether any of the macroeconomic variables can be excluded entirely from the equilibrium components of the models in exchange rate regimes 1 and 4. This is accomplished by testing zero restrictions on the full rows of the cointegrating matrix β. Our goal here is to obtain parsimonious models for the two regimes in which no variables are insignificant.

One of the problems in conducting this reduction analysis is that I(1) regressors tend to be collinear. This can lead to a situation in which a parameter estimate is initially insignificant when estimating the full composite model, but then becomes significant when one or more explanatory variables are dropped. The procedure we follow in order to address this problem is the following. First we test the exclusion restriction for each of the variables of the full composite models in exchange rate regimes 1 and 4. Second, in each regime we delete the variable with the highest p-value, and then retest the exclusion restrictions on all remaining variables. If in this second pass there are insignificant variables, then we again drop the variable with the highest p-value. We continue this procedure until all variables of the models are found to be significant. As a last step, we add back any deleted variables from the parsimonious models one at a time so as to check whether any of the variables deleted at the beginning of the process might not be significant once the parsimonious models are obtained. In no case were these deleted variables significant.

Table 4 reports the results of testing exclusion restrictions on β for the full composite model (column C) and for the parsimonious model (column P) in the two exchange rate regimes. The table shows that the data are consistent with the exclusion of relative cumulative trade balances from the model for exchange rate regime 1 and relative money supplies and U.S. secular inflation rates from the model for exchange rate regime 4. The problem of multicollinearity can be seen in the results for exchange rate regime 4, where once relative money supplies are dropped from the model, German interest rates and U.S. income and interest rates, while initially insignificant, become significant. It should be emphasized that irrespective of whether we rely on the results from the full composite models or the parsimonious models, we are led to the same conclusion, namely different sets of macroeconomic fundamentals are significant during different subperiods of floating rates.[20]

In summary, the structural change findings and the results of the cointegration and reduction analyses indicate that the process governing exchange rate movements switches from one subperiod of floating to another. They also suggest that macroeconomic fundamentals matter for the DM/$ exchange rate in some subperiods, although they matter in a striking way: different sets of macroeconomic fundamentals matter during different subperiods of the modern period of floating rates.

3 Does the Monetary Model with RE Match Up?

The finding of episodic structural change involving different sets of significant macroeconomic fundamentals during different time periods may go a long way towards explaining why empirical researchers have had such a difficult time in finding a connection between macroeconomic fundamentals and the exchange rate. Since temporal instability appears to be an inescapable feature of the empirical record, one would expect precisely what the literature finds, i.e., when one attempts to fit models with fixed coefficients over prolonged periods of floating, sometimes they fit well in sample and sometimes they do not, sometimes they forecast well and at other times their forecasting performance deteriorates markedly. The difficulty, therefore, appears to lie in the research design of

[20] It should also be noted that using a different methodology for the reduction analysis (i.e., the systems approach of Phillips [1991]) leads to the same conclusion concerning different sets of significant variables during different time periods (see Goldberg and Frydman [1996a])

virtually all empirical studies of the exchange rate, which invariably estimate exchange rate models with fixed coefficients.

The question arises as to whether the monetary model can be reconciled with the switching nature of the exchange rate process. In this section we examine this question within the context of the monetary model with RE and find that except for some highly implausible scenarios this class of models cannot be reconciled with the structural change findings. In Goldberg and Frydman [2001a] and Frydman and Goldberg [2002] we also examine the TCE framework and show that the monetary model with TCE does offer an explanation of the finding that different variables matter for the exchange rate during different time periods.

Consider the following standard monetary model due to Dornbusch [1976] and Frankel [1979]:

$$m = p + \phi y - \lambda i \tag{3}$$

$$m^* = p^* + \phi^* y^* - \lambda^* i^* \tag{4}$$

$$\dot{p} = \delta \left[\frac{\alpha}{2}(s + p^* - p - q_n) - \nu (i - i_n) \right] + \dot{\bar{p}} \tag{5}$$

$$\dot{p}^* = -\delta \left[\frac{\alpha}{2}(s + p^* - p - q_n) - \nu (i^* - i_n^*) \right] + \dot{\bar{p}}^* \tag{6}$$

$$E(\dot{s}) = i - i^* \tag{7}$$

where m, p, and y denote the log levels of domestic money supply, price and income respectively, i is the domestic short-term interest rate, q_n is the log of the natural long-run level of the real exchange rate and assumed to be constant (if absolute PPP holds, then $q_n = 0$), s is the log level of the exchange rate (defined as the domestic currency price of foreign currency), i_n denotes the natural level of domestic interest rates, defined to be the interest rate level such that when $q = q_n$, the relative excess demand for goods is zero, the symbols "*" and "·" denote foreign-country variable and time derivative respectively and $E(\dot{s})$ denotes the conditional forecast of $\dot{s}$. Equations (3) through (7) are well known. The only atypical feature is that except for the money supply variables, the parameters of the money demand functions are specified without assuming the equality of domestic parameters with their foreign counterparts. This allows the reduced-form exchange rate equation in equation (1) to be specified without the usual symmetry restrictions.[21]

The solution of this model is usually obtained by assuming rational expectations, stability and constant growth rates for domestic and foreign money and income. With these assumptions, the reduced form of the model can be written as follows:

$$E(\dot{s}) = \theta\,(\bar{s} - s) + \dot{\bar{s}} \tag{8}$$

$$\bar{s} = \ (m - \phi y + \lambda \bar{r}) - (m^* - \phi^* y^* + \lambda^* \bar{r}^*) + \lambda\pi - \lambda^*\pi^* + q_n, \tag{9}$$

where $\pi = (\dot{m} - \phi\dot{y})$ denotes expected rate of secular inflation, r denotes the domestic real interest rate and an overstrike "$-$" denotes a value associated with the instantaneous adjustment of goods prices.

Equation (9) follows directly from PPP and equilibria in the domestic and foreign money markets and is one of the cointegrating vectors of the model. The monetary model with RE also gives rise to a second cointegrating relationship, one associated with equilibrium in the foreign exchange market and that follows from uncovered interest rate parity, RE and flexible prices:

$$i - i^* = \pi - \pi^* = \bar{i} - \bar{i}^* \tag{10}$$

Of course along the saddle path, deviations of nominal interest rates from their goods-market equilibrium values are uniquely tied to the deviations of the exchange rate from its goods-market equilibrium value.[22] Thus, the monetary model with RE, stability and constant growth rates of money and income implies only one linearly independent cointegrating relationship, i.e., both cointergrating vectors in equations (9) and (10) imply equilibria in the goods markets and the foreign exchange market. Since the cointegrating VAR setup of Johansen [1988] estimates linearly independent cointegrating vectors, the standard solution of the monetary model with RE gives rise to a Π matrix with a rank of one. Of course, money and income may not be exogenous to the model, so that the

[21] The assumption of a constant q_n has no bearing on the implications of TCE. In the empirical applications of the preceding section we relaxed this assumption along the lines of Hooper and Morton [1982] and modeled q_n as a function of relative cumulative trade balances. Note also, that Frankel [1979] excludes short-run interest rates from the price-adjustment equations in (5) and (6). Goldberg [1995] shows that the inclusion of these variables is necessary for monetary policy actions to be neutral in the long-run *with RE.*

[22] Namely, $s - \bar{s} = -\frac{1}{\theta}\left[(i - i^*) - \left(\bar{i} - \bar{i}^*\right)\right]$, where θ denotes the stable root of the system, redefined to be positive.

monetary model with RE is potentially consistent with our finding of multiple cointegrating vectors in exchange rate regimes 1 and 4.[23]

Given the foregoing model, there are just two possibilities for different sets of macroeconomic fundamentals to matter for exchange rates during different time periods. Either this kind of structural change occurs through the semi-reduced form of the model (i.e., in equations (3) through (6)) and/or it occurs through the expectations channel. The first possibility implies the underlying structure of the economy undergoes structural change of a radical nature, so for example, sometimes income is important for money demand considerations and sometimes it is not. Although there are many studies documenting the instability of money demand functions, there is nothing in the literature suggesting that money demand functions experience the kind of radical structural change that would be required to explain the structural change findings of the preceding section.[24]

As for the second possibility, the RE approach does allow for different variables to matter for forecasting during different time periods, but only if such changes occur in the underlying structure of the economy. It is tempting to rely on the Lucas critique, but as we show in Frydman and Goldberg [2001b], changes in policy reaction functions do not lead to different sets of macroeconomic fundamentals coming into the cointegrating exchange rate relation within the context of the monetary model. The problem is that with RE, there is a rigid connection between agents' forecasting functions and the underlying structure of the economy, so that the kind of episodic structural change found in this paper can be explained only as a result of structural change in the semi-reduced form. There is no room with RE for market agents to believe that cumulative trade balances are relevant for short-run exchange rate movements during some time periods and not during others. If the one true model indicates the importance of cumulative trade balances, then it becomes irrational to ignore this factor in forming expectations. Thus, unless one is prepared to argue that the underlying structure of the economy experiences episodic structural change of a radical nature, the monetary model with RE is inconsistent with our structural change findings.

4 Conclusion

The findings of this study lead to three conclusions. First, the process governing the short-run movement of exchange rates over periods of floating is episodically

[23] It is interesting to note that there are a number of studies that attempt to estimate PPP and UIP relations using the Johansen setup and find evidence for PPP, but only when the UIP relation is present (e.g., Johansen and Juselius [1992], Juselius [1995] and Diamandis, Georgoutsos and Kouretas [2000]). But since the Johansen setup distinguishes only linearly independent cointegrating vectors, the finding of both UIP and PPP is problematic for the standrad monetary model with RE. On this point see Goldberg [2001].

[24] For a recent study on money demand see Goldfeld and Sichel [2000].

unstable. Second, the switching nature of the exchange rate process takes a striking form, namely that different sets of macroeconomic fundamentals matter for exchange rates during different time periods. Finally, it is difficult to reconcile these structural change findings with the monetary model with RE.

The implausibility of relying on changes in the structure of the model under RE as a way to explain the temporal instability of the monetary model suggests that what is required is an alternative expectational framework that breaks the rigid connection between expectations functions and the one "true" semi-reduced form. The TCE framework achieves this end by recognizing that agents possess a pluralism of theories with which to look forward and it assumes that their theories provide only qualitative knowledge about how the exchange rate might be related to some set of fundamentals.[25] With imperfect knowledge, as formulated by the TCE framework, agents may change the theories (or sets of fundamentals) they use to forecast exchange rate movements from one time period to another. It is this possibility of changing beliefs that allows the monetary model with TCE to be reconciled with the our finding that different sets of macroeconomic fundamentals matter for short-run exchange rate movements during different time periods. The conclusion is that if one is prepared to move away from the RE paradigm and recognize that market agents possess imperfect knowledge, then plausible interpretations of the connection between floating exchange rates and macroeconomic fundamentals might after all become possible.

[25] Theories used by agents are assumed to inform them in possibly three qualitative ways: 1) they may inform agents as to the important explanatory variables that should be included in their forecasting equations; 2) they may indicate the algebraic signs of the weights that should be attached to these fundamental variables; and 3) they may say something about long-run equilibrium levels. Expectations functions possessing explanatory variables matching those implied by agents' theories and where the weights attached to these explanatory variables are consistent in sign with one or more of these theories are said to be theories consistent.

Data Appendix

All data are monthly. The data set begins in March 1973 and ends in November 1998. The trade balance and nominal GDP data are from the O.E.C.D. (M.E.I.) data bank. All other time series are from the IFS data bank.

- – s — end of month, DM/$
  - i — end of month 3-month treasury bill rate
  - i* — end of month 3-month interbank deposit rate
  - π and π* — proxied by using the average inflation rate (based on the CPI) over the preceding 12 months.
  - y and y* — an index of industrial production, seasonally adjusted
  - m and m* — M1, end of month, in billions of local currency
  - k_r — domestic and foreign trade balances are first normalized by dividing by the 1985 monthly average of nominal GDP, and then the relative magnitude was cumulated

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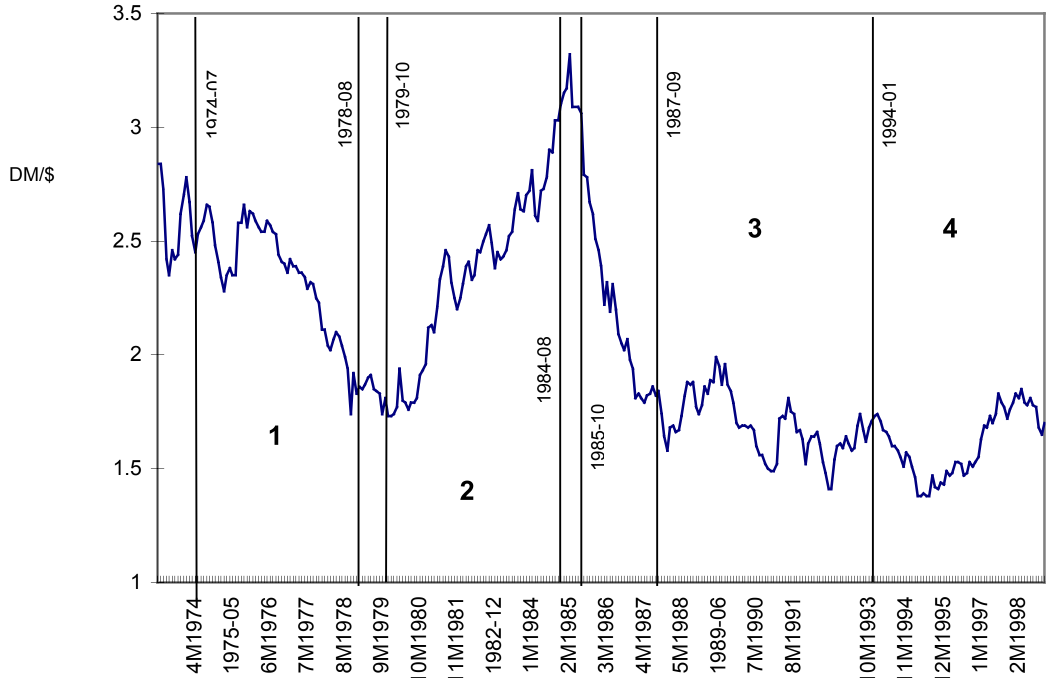
Figure 1



DM/$
3.5
3
2.5
2
1.5
1
1974-07
1978-08
1979-10
1984-08
1985-10
1987-09
1994-01
1
2
3
4
4M1974
1975-05
6M1976
7M1977
8M1978
9M1979
10M1980
11M1981
1982-12
1M1984
2M1985
3M1986
4M1987
5M1988
1989-06
7M1990
8M1991
10M1993
11M1994
12M1995
1M1997
2M1998

Table 1
Diagnostic Tests of the Unrestricted VAR[a,b]

Equation	Time Period and Order of System						
	73:03-98:12	74:07-78:09		79:10-84:08	87:09-93:12[c]	94:01-98:12	
	AR(8)	AR(1)		AR(3)	(AR(3)	AR(1)	
	C	C	P^c	C	C	C	P^e
s		A^{*}	H^{*}, N^{***}	S^{***}	S^{**}	N^{*}	
m_r	N^{***}, S^{***}	S^{*}	S^{*}		N^{***}, S^{***}		
y	N^{***}			S^{***}	S^{***}		
y*	N^{***}, S^{*}				S^{**}		
π	N^{***}				N^{***}		
π*	A^{**}, N^{***}			S^{***}			
i	A^{***}, H^{**} N^{***}, S^{***}		A^{*}		S^{*}		
i*	A^{***}, H^{*} N^{***}	A^{*},N^{**}	H^{*}, N^{***}	S^{***}	S^{*}	H^{***}, N^{***}	H^{***}, N^{***}
k_r	A^{***}, N^{***} S^{***}	S^{**}				S^{*}	S^{*}

a. Table reports significance levels of four tests: 1) The F-form of the ARCH test (A); 2) The White (1980) heteroscedasticity test (H); 3) The Doornik and Hansen (1994) chi square test for normality (N); and 4) An F-statistic for serial correlation (S). Significance levels are indicated, where at .01, .05 and .10 levels by ***, **, * respectively.
b. Columns labeled C denote test statistics for the full composite model and columns labeled P denote test statistics for the parsimonious models.
c. A dummy for money supply running from 1990:06 to 1991:01 was included in the estimation.
d. The parsimonious model excludes k_r.
e. The parsimonious model excludes m_r and π*.

Table 2
Testing for the Number of Cointegrating Vectors
Adjusted TR and MAX Statistics[a]

Number of Co-integrating Vectors	Exchange Rate Regime[b]							
	1974:07-1978:09				1994:01-1998:11			
	C		P[c]		C		P[d]	
	MAX	TR	MAX	TR	MAX	TR	MAX	TR
$r \leq 0$	**58.80***	**260.20*****	**60.06****	**220.00*****	**80.04****	**300.60*****	**82.25*****	**239.90*****
$r \leq 1$	51.15	**201.40*****	**46.73***	**160.00****	**53.19***	**220.60*****	**52.55*****	**157.60*****
$r \leq 2$	43.28	**150.20****	31.68	**113.20***	**48.43***	**167.40*****	**39.87****	**105.10*****
$r \leq 3$	32.17	106.90	26.22	81.56	36.91	**119.00****	**28.89***	**65.22****
$r \leq 4$	24.10	74.77	20.69	55.34	34.12	82.06	19.33	36.33
$r \leq 5$	19.59	50.67	18.31	34.65	17.16	46.84	11.46	17.00
$r \leq 6$	15.19	31.08	10.48	16.33	16.25	29.78	5.53	5.53
$r \leq 7$	10.20	15.90	5.86	5.86	8.06	13.53		
$r \leq 8$	5.70	5.70			5.48	5.48		

a. TR and MAX denote the trace and max eigenvalue statistics of Johanson (1988). Significance levels are based on Osterwald-Lenum (1992). Sinificance at the .01, 05, and .10 and indicated by ***, *** and * respectively.
b. Columns labeled C denote rank tests for the full composite model and column labeled P denote the rank tests for the parsimonious models.
c. The parsimonious model excludes k_r.
d. The parsimonious model excludes m_r and π*.

Table 3
Testing the Exchange Rate for Weak Exogeneity[a]

Exchange Rate Regime			
1974:07-1978:09		1994:01-1998:11	
C	P[b]	C	P[c]
10.12** (.018)	**7.87**** (.048)	**8.61*** (.072)	**10.39**** (.016)

a. Tables reports of zero restrictions on the first (exchange rate) row of the α matrix, i.e., $\alpha_{1j} = 0$ for all j, where j = 3 for exchange rate regime 1 and j = 4 for exchange rate regime 4. The colums labeled C denote test statistics for the full composite model and the columns labeled P denote test statistics for the parsimonious model. Significance at the .05 and .10 is indicated by ** and * respectively.
b. The parsimonious model excludes k_r.
c. The parsimonious model excludes m_r and π^*.

Table 4
Testing Exclusion Restrictions on β[a]

Variables	Exchange Rate Regime			
	1974:07-1978:09		1994:01-1998:11	
	C	P	C	P
s	10.04 (.018)	11.51 (.009)	13.24 (.010)	24.50 (.000)
m_r	6.40 (.093)	7.30 (.063)	**4.30**# (.367)	–
y	16.08 (.001)	21.66 (.000)	13.81 (.008)	29.50 (.000)
I	8.11 (.044)	9.94 (.019)	**6.12**# (.191)	8.41 (.078)
π	7.01 (.072)	7.62 (.055)	29.20 (.000)	43.67 (.000)
y*	19.44 (.000)	21.09 (.000)	**7.75**# (.101)	14.89 (.005)
i*	8.12 (.044)	7.86 (.049)	**7.20**# (.126)	11.50 (.021)
π*	8.02 (.046)	19.00 (.000)	**5.17**# (.270)	–
k_r	**5.36**# (.147)	–	8.05 (.090)	16.55 (.002)

a. Table reports on zero restrictions on the full row of the β matrix for each variable using the likelihood ratio statistics of Johansen (1988) and Johansen and Juselius (1990). P values are given in parentheses. The symbol A „-" denotes that the variable has been dropped from the model due to insignificance. The symbol # denotes insignificance with a p-value greater than .10.

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1) vergriffen (out of print)
2) In abgeänderter Form erschienen in Berichte und Studien Nr. 4/1990, S 74 ff
3) In abgeänderter Form erschienen in Berichte und Studien Nr. 4/1991, S 44 ff
4) In abgeänderter Form erschienen in Berichte und Studien Nr. 3/1991, S 39 ff
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