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Empirical Exchange Rate Models and Shifts in the Co-Integrating Vector*

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This paper finds that many of the anomalies in the exchange rate literature can be explained in large measure by recurrent shifts in the co-integrating vector. Using structural change tests that search recursively for break points, we find that the co-integrating vector implied by a composite monetary model experiences parameter shifts on five occasions over a sample period that includes the 1970s and 1980s. Within the implied regimes of parameter constancy, we find that many of the fundamental variables of the monetary models are significant and enter with correct parameter signs. We also find that the sets of significant variables in each regime are cointegrated, implying the existence of stable long-run relationships operating in the foreign exchange market. Finally, all of the structural models examined outperformed the random walk model in out-of-sample fit by considerable margins within the separate regimes of stability (in some cases by a margin of 70 percent in root mean square error), indicating that the large forecasting errors reported in Meese and Rogoff [1983] are the result of allowing the forecasting experiment to run past the end of one exchange rate regime and into the next. Our findings also suggest that a primary factor behind the observed instability of the co-integrating vector is unstable exchange rate expectations functions.

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Empirical Exchange Rate Models and Shifts in the Co-Integrating Vector

There is a widespread belief that standard empirical exchange rate models provide little towards explaining the modern period of floating exchange rates. The empirical record includes the following: 1) an inability on the part of structural models to outperform the random walk model when based on out-of-sample fit (e.g., Meese and Rogoff [1983]); 2) a lack of co-integrating relationships among the exchange rate and set of supposed fundamentals (e.g., Meese and Rogoff [1988] and Meese and Rose [1991]); and 3) parameter estimates that are either insignificant or significant and of the wrong sign (e.g., Dornbusch [1980], Frankel [1983], and Boothe and Glassman [1987]).

The purpose of this paper is to investigate whether the failure of standard empirical exchange rate models is due in part to periodic shifts in the co-integrating vector. Standard exchange rate models all posit a long-run relationship between the exchange rate and some set of macroeconomic fundamentals, which is usually expressed in a log-linear reduced form. Since the exchange rate and most macroeconomic fundamentals are commonly viewed as being integrated of order one (I(1)), then the equilibrium assumptions of standard exchange rate models imply that their reduced forms involve at least one vector of parameter values which is co-integrating.[1] However, if the long-run relationship between the exchange rate and macroeconomic fundamentals undergoes shifts from time to time, then no one vector of parameter values will be co-integrating. Instead, each stable long-run

[1]As such, the current practice of working with exchange rate equations which are specified in first differences (e.g., Meese [1986], Meese and Rogoff [1988] and Papell [1988], among many others) is inappropriate for testing the validity of standard exchange rate models. Although such a practice circumvents the problem of spurious regressions, it implies the absence of a long-run relationship, the very claim made by standard theory. If a relationship is found in first differences and (by construction) not in levels, then a new theory of exchange rate dynamics is required.

relationship will give rise to a time period characterized by a different co-integrating vector. From an empirical viewpoint, we would expect in such a case that a series of successive linear reduced-form equations (i.e., a piece-wise linear specification) would provide a reasonable approximation to the data.

There are several reasons extant in the literature suggesting that this may in fact be the case. From a theoretical standpoint, the Lucas critique indicates that periodic changes in government policy may lead to periodic changes in exchange rate expectations functions and, therefore, to periodic changes in reduced-form parameters and the co-integrating relationship. Furthermore, there are a number of recent studies suggesting that expectations functions may undergo periodic shifts independently of policy changes. This may occur because trading horizons are short (Froot, Scharfstein and Stein [1992]), or because market agents are endowed with only qualitative knowledge of the model governing exchange rate movements, which we call theories consistent expectations (Goldberg and Frydman [1993a,b]). These studies show that market agents may focus on only a subset of the relevant fundamental variables and this subset may change from one time period to the next.

From an empirical standpoint, it is widely known that parameter instability is an important problem for empirical exchange rate models. The studies of Boughton [1987] and Meese and Rogoff [1988], among others, find structural breaks which are proximate to the early Volcker years of tight U.S. monetary policy.[2] 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)."

[2]Note, Meese and Rogoff [1988] test for structural change using a model specified in first differences.

Furthermore, the studies of Engel and Hamilton [1990] and Kaminsky [1993] indicate that the univariate process governing exchange rate movements does in fact undergo discrete switches. Although these studies are confined to a univariate analysis, their results are suggestive that the multivariate process may also exhibit such switching behavior.

Yet, despite this evidence suggesting that the long-run (multivariate) relationship between the exchange rate and macroeconomic fundamentals is and should be periodically unstable, the evidence used in illustrating the failure of standard exchange rate models largely ignores this problem. Invariably, empirical studies on exchange rates use sample periods covering the 1970s and 1980s in estimating co-integrating vectors and testing for in-sample and out-of-sample fit.[3]

In this paper we address the problem of periodic shifts in the co-integrating vector by first testing a composite monetary model of the U.S. dollar/German mark exchange rate for discrete points of parameter instability over a sample period that includes the 1970s and 1980s. Data mining is avoided by using structural change tests that search recursively for break points, rather than using tests that impose break points a priori. We use the structural change results to distinguish subperiods or regimes characterized by *relative* parameter constancy and, therefore, by a stable long-run relationship. We then test several standard monetary models within each regime for cointegration and in-sample and out-of-sample fit.

There are studies that confront the unstable nature of empirical exchange rate models with random coefficient techniques and they find some improvement in out-of-sample fit (Alexander and Thomas [1987] and Schinasi and Swamy [1989]). It is important to point out that random coefficient techniques are not designed to capture discrete changes in parameter values, but rather stochastic

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[3]The term "long-run relationship" must be used with care. In a world where the long-run or co-integrating relationship shifts periodically, the system may never reach the long-run. Yet, in any one time period, there is a long-run relationship governing exchange rate dynamics.

variation of parameter values around fixed means. As such, they are not capable of capturing periodic shifts in the co-integrating vector. Although some random variation of parameter values most likely occurs in all time periods (hence the phrase "regimes of relative parameter constancy"), our aim in this paper is to examine whether discrete changes in the co-integrating vector are perhaps a more important factor for the fit of empirical exchange rate models.

There are also studies that attempt to assess the importance of non-linearities for empirical exchange rate models and they conclude that there is no evidence of such phenomena (Meese and Rose [1991] and Chinn [1991]). It is important to point out, however, that the techniques employed in these studies are not capable of capturing the type of piece-wise linearity that arises due to the temporal instability of the co-integrating relationship. The basic difference between these studies and ours is that we contemplate non-linearity with respect to time, whereas the former search for non-linearity according to other metrics, such as the magnitudes of the variables.[4]

Our findings can be summarized as follows: 1) the co-integrating vectors implied by standard monetary models of the exchange rate experience parameter shifts on more occasions over the modern floating rate period than previously documented, giving rise to three periods of relative parameter stability; 2) many of the fundamental variables of standard theory are significant and enter with correct parameter signs within the separate regimes of stability; 3) different sets of fundamental variables are significant during different time periods; 4) these sets of significant variables are cointegrated, implying the existence of stable long-run relationships operating in the foreign exchange market; and 5) all of the structural models examined outperform the random walk model in out-of-

[4]We do not deny that the type of non-linearities considered in Meese and Rose [1991] and Chinn [1991] may be important. However, our results suggest that in testing for such non-linearities it may be important to first locate subperiods involving stable long-run relationships. The studies of Meese and Rose [1991] and Chinn [1991] ignore this problem and apply nearest neighbor and other methods to sample periods covering the 1970s and 1980s.

sample fit by considerable margins within the separate regimes of stability (in some cases by a margin of 70 percent in root mean square error), indicating that the large forecasting errors reported in Meese and Rogoff [1983] are the result of allowing the forecasting experiment to run past the end of one exchange rate regime and into the next. Hence, we find that not only do periodic shifts in the co-integrating vector occur, but that such shifts explain several of the anomalies extant in the literature.[5]

Our findings raise two important questions. First, what factors lie behind the observed instability of the co-integrating vector? And second, can any of the standard monetary models of the exchange rate be reconciled with our findings. As for the first question, our finding that different sets of fundamental variables matter during different time periods is not only suggestive that shifts in expectations functions are an important factor, but it lends support to the alternative approaches of Froot, Scharfstein and Stein [1992] and Goldberg and Frydman [1993a,b]. We also find that many of the observed points of parameter instability match up well with time periods involving major changes in government policy and this also points to the importance of the expectations channel. As for the second question, we attempt a detailed answer in Goldberg and Frydman [1991], where we find that a sticky-price monetary model with theories consistent expectations is consistent with our empirical findings.

The structure of the paper is as follows. Section 2 presents the methods and results of the structural change analysis, while section 3 presents the methods and results of the regression and co-

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[5]It is important to emphasize that the Meese and Rogoff analysis is an exercise in out-of-sample fit and not in forecasting. As such, it is appropriate not only to use the actual future values of the explanatory variables (as in Meese and Rogoff), but also to use the structural change results. This is not to say that forecastability is unimportant for assessing the merits of empirical exchange rate models. Our results suggest that such an analysis, which we do not attempt in this paper, would require the added dimension of forecasting future break points.

integration analysis. Section 4 reports on the results of testing for out-of-sample fit. In section 5 we offer concluding remarks.

2.0 Distinguishing Regimes of Relative Parameter Stability

In carrying out the analysis of this and the other sections we use monthly data for the U.S. dollar/German mark exchange rate over a period that begins in March 1973 and ends in March 1988.[6] We also use a composite monetary model of the exchange rate developed in Goldberg [1993a,b] called the Dornbusch et. al. (D+) model. The D+ model encompasses three of the more frequently used models in both the theoretical and empirical literature: 1) the Frenkel [1976] and Bilson [1978,1979] (FB) flexible-price model; 2) the Dornbusch [1976] and Frankel [1979] (DF) sticky-price model; and 3) the Hooper and Morton [1982] (HM) sticky-price model. In addition, the D+ model gives rise to a reduced form in which the explanatory variables do not enter with the usual restriction that parameters on domestic variables equal in absolute terms their foreign counterparts.[7] The reduced form is as follows:

$$e = \beta_0 + \beta_1 m + \beta_2 m^* + \beta_3 y + \beta_4 y^* + \beta_5 i + \beta_6 i^* + \beta_7 \pi + \beta_8 \pi^* + \beta_9 k + \beta_{10} k^* , \qquad (1)$$

where m, p, and y denote the log levels of domestic money supply, price and income respectively, i denotes the level of the domestic short-term interest rate, π is the expected long-run inflation rate,

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[6]For further details on the data set see the appendix.

[7]In order to obtain such a reduced form, two extensions to the standard setup are required. First, the standard money demand restrictions setting parameters equal across countries must be dropped and second, the assumption of perfect capital mobility must be relaxed along the lines of Frenkel and Rodriguez [1982]. For the solution and theoretical discussion of such a model see Goldberg [1993a,b]. It should be noted that the D+ model does not assume imperfect capital mobility, but rather allows for this possibility.

k is the de-trended level of domestic cumulative trade balances and a "*" denotes foreign country variable. Except for the distinct parameter values, equation (1) is well known. The predictions of the various monetary models as to the signs of the β-coefficients are provided in table 1.[8]

In order to test the empirical exchange rate models subsumed in equation (1) for periodic shifts in the co-integrating vector we make use of two procedures, the cusum test and the Quandt ratio (QR) technique as given in Brown, Durbin and Evans [1975]. The cusum test is used to establish (in a statistical sense) that a break has occurred and the QR technique is used to determine the most likely location of the break point.[9] It should be emphasized that these procedures search the data recursively for the possibility of one or more break points, rather than relying on tests that require the choosing of break points a priori. Thus, they provide *a mechanical, non-discretionary* way for locating points of parameter instability, thereby avoiding data mining. Another feature of the cusum test is that is valid for a model involving I(1) variables. All of the variables contained in equation (1) are generally recognized to be I(1).[10] The problems with such nonstationary variables are now well known (e.g., Granger and Newbold [1974] and Phillips [1986]). Unlike F-stat based tests, the cusum test can be applied without complication because it is based on standardized recursive OLS residuals, which with I(1) variables are super-consistent (Stock [1987]).

In order to test for the possibility of a number of break points in the co-integrating vector we

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[8]Note, the parameters on both interest rate variables may be positive or negative depending on the degree of capital mobility. For a discussion of this result see Goldberg [1993b].

[9]This procedure of first using the cusum test to determine that a break has occurred and then the QR technique to determine its approximate location is suggested in Frydman [1983]. Boughton [1987] also uses a similar procedure.

[10]For example, see Meese [1986] and Meese and Rogoff [1988]). In Goldberg [1991] the $/DM exchange rate and explanatory variables of equation (1) are subjected to a large battery of unit root tests including the ADF tests of Fuller [1976] and Dickey and Fuller [1981] and the Z tests of Phillips [1987] and Perron [1988]. In every case the null of a unit root cannot be rejected.

employ the following strategy. We start at the beginning of our sample in March 1973 and test the D+ model for parameter instability using the cusum test and QR technique. Once a break point is found, we reduce the composite model (if possible) by deleting those variables found to be statistically insignificant within the implied region of relative parameter constancy.[11] We also check the residuals of this reduced model for serial correlation.[12] In order to reduce the influence of this problem if detected, we add the following regressors to the regression equation: 1) first differences of each of the explanatory variables; 2) two lags of each of these first difference terms; and 3) two lags of the exchange rate first differenced. We find that the inclusion of these stationary I(0) variables, which does not disturb the consistency of the parameter estimates on the I(1) explanatory variables and hence the consistency of the residuals (see Granger [1986]), is able to account for most of the serial correlation present when found.[13] We then re-test for instability using the reduced model. Once the first break point is established, we re-run the cusum test and QR technique using the D+ model, but with a subsample that begins after the location of the first break point.[14]

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[11]It is important to test for instability using the reduced model because the presence of insignificant I(1) regressors may bias the test results. This is because the problem of multicollinearity frequently arises with I(1) variables. See Hendry [1986] for a discussion of this issue. The methods by which we obtain reduced models are discussed in the next section.

[12]The problem with serially correlated residuals is that they can distort the size of the cusum test, since the confidence bounds are constructed with the assumption that the cumulative sum of standardized recursive residuals involves independent steps under the null hypothesis.

[13]Since the variables of the reduced models are found to be cointegrated, the addition of the I(0) terms can be viewed as capturing the short run dynamics of the system. It is true that structural breaks found with the augmented models may be due in part to shifts in the parameters of the first-difference terms. However, as the next section makes clear, there is no doubt that the parameters of the static long run relationship are also shifting.

[14]It is necessary to restart the structural change procedures once a break has been detected because after this point the cusum no longer has an expected value of zero, whether or not the true model returns to the original relationship.

Additional break points are found in a similar way.

It should be pointed out that the location of the break points obtained with the composite model and the reduced models match up in every case except one (see footnote 17). As such, the problems of multicollinearity and serially correlated residuals are found to be of little significance for the results of the structural change analysis.

2.1 The Structural Change Results

The results of our structural change analysis are summarized in figure 1, where break points are indicated by dotted vertical lines. The figure indicates that a total of five break points are found, which in turn gives rise to three subperiods of relative parameter stability, i.e., three regimes characterized by relatively stable long-run relationships (the unshaded regions numbered 1 through 3).[15] The first exchange rate regime begins in July 1974 and runs through September 1978, the second begins in November 1979 and runs through August 1984 and the last begins in September 1985 and runs until the end of the sample in March 1988.

There are several noteworthy aspects about the results in figure 1. First, after the first two exchange rate regimes and after the collapse of the Bretton Woods era of fixed rates there are periods of transition, i.e., periods of time which are relatively transient and which are distinct from the regime they immediately follow (these are the shaded regions in figure 1). The reasons for why they arise may include uncertainty on the part of market agents as to which model is currently relevant (Goldberg and Frydman [1993a,b]), the choosing of a model not encompassed by the D+ model, the occurrence of a (rational or irrational) stochastic bandwagon movement (e.g., Evans [1986] and Meese [1986]), or a movement away from relying on a model based on fundamentals to one based

[15]The actual cusum and QR plots are reported in Goldberg [1991].

on chartism (e.g., Frankel and Froot [1987] and Kirman [1990]).[16]

The transition period following the Bretton Woods era is of particular interest because it provides justification for the common practice in the literature of deleting some arbitrary period of time at the beginning of the floating rate period prior to analysis (e.g., see Frankel [1979,1983] and Boothe and Glassman [1987], among many others).[17] This practice is usually based on the belief that the market required some amount of time in order to acquaint itself with the new regime of floating rates. The period following this interval of learning is dubbed the period of mature floating. Our structural change results show that not only is the long-run relationship governing exchange rate movements different in this first subperiod of floating (thus validating the story of immature floating), but they also provide objective criteria for determining its duration.

Another interesting aspect of the results in figure 1 is that each of the break points after the break in July 1974 (which ushers in the period of mature floating) corresponds with a major shift in policy. The break in October 1978 (break point A in figure 1) was preceded by a protracted period of dollar depreciation and accelerating U.S. inflation. This in turn led to very low real

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[16]The break point in November 1979, which gives rise to the transition period following the first exchange rate regime, is not based on the cusum and QR tests, but rather is based on a Chow test using the systems approach of Phillips [1991]. (This systems approach, which we discuss more fully in the next section, allows for valid inference to be drawn from an F-statistic.) Ideally, the structural change tests should be able to locate such transition periods if they exist. However, these periods may be too transient for the cusum test to detect. In light of this problem and the fact that transition periods are found after the Bretton Woods era of fixed rates and after the second exchange rate regime, we check for a structural break in November 1979. November 1979 is chosen because it is the period in which the value of dollar finally turns around. The systems based Chow test indicates a break at the 2.5 percent significance level.

[17]The model used in locating the break point in July 1974 is the D+ model without a constant. When a constant is added to the regression, the cusum and QR tests indicate that a break does not occur until October 1978. However, when a reduced model is fitted to this first (possible) region of stability (i.e., from March 1973 through September 1978) the constant term is found to be insignificant.

interest rates in the U.S. and continued capital outflow. In order to stem the troubling tide the U.S. implemented a package of measures in November 1978 (which was perhaps anticipated in October 1978), which included the Carter bond program of active intervention (the level of intervention for the month of November in support of the dollar was $6.86 billion, a record at that time) and measures to tighten monetary policy. In addition, the German Bundesbank was also intervening at very high levels (see Frankel [1982]). These measures were initially successful as the dollar firmed in the last quarter of 1978. However, in the second quarter of 1979 the dollar began its slide once again. Volker was appointed in August 1979 and, in October of that year (break point B in figure 1), the Fed announced that it was de-emphasizing the federal funds rate as an operating target. Instead, the primary operating target became nonborrowed reserves. This shift in monetary policy, with its associated rise in U.S. interest rates, was enough to finally turn around the trend in the dollar.

In terms of the 1980s, the dollar and U.S. trade deficit had climbed so high by the first quarter of 1984 that the stated Reagan (and Regan) policy of laissez faire was changed to one of actively trying to talk the dollar down. Intervention picked up after a long absence and with the 1984 presidential election over, there was a major change in policy (perhaps again anticipated in September 1984, which is break point C in figure 1). James Baker took over as Treasury Secretary in January 1985 with the stated goal of getting the dollar down and improving the trade picture. He engineered an international agreement that involved reducing the U.S. budget deficit and relaxing U.S. monetary policy (Volcker agreed to cooperate) in return for fiscal stimulus and high interest rates abroad. Although this policy of international coordination was discussed at the Bonn economic summit in May 1985, it was not formally announced by the G-7 until September 1985 (break point D in figure 1) at the Plaza hotel in New York.

The fact that many of the break points in figure 1 line up with periods involving major changes in government policies is suggestive that the structural change results are due in part to periodic shifts in exchange rate expectations functions, either via the Lucas critique or via the channels postulated in Froot, Scharfstein and Stein [1991] and Goldberg and Frydman [1993a,b]. In the next section we provide additional evidence supporting the view that periodic shifts in expectations functions lie at the heart of the instability findings.[18]

3.0 Regression Analysis and Tests of Cointegration

Much of the evidence on the asset market models from in-sample regression studies is based on using standard techniques such as OLS and linear reduced forms specified in log-levels (e.g., Frankel [1979,1983] and Boothe and Glassman [1987]). But, with I(1) variables, such an approach is inappropriate due to the problem of spurious regressions (Granger and Newbold [1974] and Phillips [1986]). It may be the case, however, that although a group of variables are individually nonstationary, there may exist market forces that work over time causing these variables to move together, i.e., there may exist some linear combination of the I(1) variables that is stationary. In such a case the variables are said to be cointegrated.[19] Testing for cointegration is important because it provides a direct way to test for the existence of a stable long-run relationship among I(1) variables, the very claim made by standard exchange rate theory (see footnote 1). Studies that test

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[18]It may appear that the Lucas critique provides a complete explanation for the instability findings. However, in the next section we find that different sets of variables are significant during different time periods. Such a finding is not only difficult to reconcile with the standard RE framework, but it is the outcome of expectations behavior explicitly modeled in Goldberg and Frydman [1991,1993a,b]. We discuss this more fully in the next section.

[19]The early articles developing the theory of cointegration include Granger [1983] and Engel and Granger [1987]. See also Hendry [1986] and the other articles contained in the special issue.

for co-integration include Meese [1986], Meese and Rogoff [1988] and Meese and Rose [1991]. All of these studies, which estimate co-integrating vectors over sample periods that include the 1970s and 1980s, are unable to reject the null hypothesis of no cointegration.

The purpose of this section is twofold. First, we make use of a systems approach recently developed in Phillips [1991] that allows us to draw valid inferences concerning cointegrating vectors. We use this systems approach in order to estimate reduced models (by deleting statistically insignificant regressors) for the first and second exchange rate regimes.[20] Second, we test the variables of our reduced models and the models listed in table 1 for cointegration.

3.1 The Method of Reduction

The systems approach of Phillips [1991], which extends the work of Johansen [1988], applies maximum likelihood (ML) estimation on a system that specifies the unit roots explicitly. This system can be expressed as follows:

$$\begin{bmatrix} y_t \\ \Delta x_t \end{bmatrix} = \begin{bmatrix} 0 & b' \\ 0 & 0 \end{bmatrix} \begin{bmatrix} y_t \\ x_t \end{bmatrix} + \begin{bmatrix} u_t \\ v_t \end{bmatrix}, \qquad (2)$$

where y_t denotes an I(1) dependent variable, x_t is a vector of I(1) explanatory variables, b is the cointegrating vector and u_t and v_t are generated by a stationary vector ARMA model. Phillips [1991] shows that when y_t and x_t are cointegrated, the ML estimator of b is consistent, unbiased and

[20]We are not able to estimate a reduced model for the third exchange rate regime due to a lack of degrees of freedom, since the systems approach of Phillips [1991] requires the estimation of 30 parameters when testing the full composite model. See section 3.1.

possesses standard asymptotics.[21]

In order operationalize this systems approach, the ARMA process underlying u_t and v_t must be specified. This raises some difficulty, however, because the D+ model involves ten explanatory variables, implying that the system in (2) involves 11 equations. In light of this and because the longest regime of stability consists of 58 observations, the analysis is limited to considering AR errors only. An AR specification is arrived at in a sequential manner using the system of equations in (2). For each exchange rate regime, we begin by estimating (2) with an AR(2). Then, all the AR terms found to be insignificant are deleted and the system in (2) is re-estimated. If the residuals of any of the 11 equations are found to be autocorrelated (by regressing these residuals on 10 lags), then additional AR terms are added until independent residuals are obtained. In all 11 cases, tests of the residuals produced by this procedure for autocorrelation could not reject the null hypothesis of no autocorrelation. As such, AR errors are found to fit the data reasonably well.

As mentioned earlier, one of the problems in working with I(1) regressors is that they may appear correlated even though there is no genuine relationship between them. This problem of collinearity can lead to a situation in which a parameter estimate is initially insignificant when estimating the composite model, but then becomes significant (or even changes sign) when one or more explanatory variables are deleted. The procedure we follow is to begin by deleting one explanatory variable at a time, beginning with the variable whose t-value is the lowest. However, the problem of multicollinearity raises legitimate questions about whether the results we obtain are

[21]An estimator of the cointegrating vector with standard asymptotic properties can also be obtained using the procedure developed in Phillips and Hansen [1988], which adjusts the OLS parameter estimates and standard errors of the cointegrating regression nonparametrically. The authors call this procedure fully modified least squares (FMLS). We initially used this procedure in estimating the composite model, but the results were found to be overly sensitive to the value of the lag truncation parameter used in estimating the long-run covariance terms. This approach was therefore found to be unusable. However, Nicoletti [1990] uses FMLS with much success.

dependent on the variables chosen to be deleted. In order to examine the sensitivity of our results we add the deleted variables back in different combinations. This provides an additional check that these variables are not deleted because of a spurious relationship among the regressors. It should be emphasized, however, that although the reduction results need to be interpreted with care, the other key results of this paper (i.e., the structural change, co-integration and out-of-sample fit results) are insensitive to the way in which we obtain reduced models.

3.2 The Regression and Co-Integration Results

Table 2 presents the results of the regression analysis. Regressions 5 and 9 provide the outcome of our reduction analysis, which are the reduced models for the first and second exchange rate regimes respectively.[22] From an empirical standpoint, these regression results show that unlike previous studies, many of the explanatory variables of the monetary models are statistically significant with correct parameter signs (see table 1).[23] From an economic viewpoint, it is not clear how much support the results in table 2 provide for any one monetary model of the exchange rate. Extensive analysis of this issue is provided in Goldberg and Frydman [1991], where we find that the results of regressions 5 and 9 are consistent with the sticky-price monetary framework with theories

[22]The results of the complete step-by-step reduction analysis followed in obtaining these reduced models are reported in Goldberg and Frydman [1991].

[23]The seemingly anomalous parameter signs for the money supply variables during the second exchange rate regime are sensible if (unanticipated) changes in money lead to changes in interest rates in the same direction. There are a number of studies which find that this is in fact the case for the U.S. during the early 1980s and conclude that this behavior is the result of credible central bank targets, as opposed to the view that changes in money cause agents to revise their inflation forecasts in the same direction (Engel and Frankel [1982,1984], Cornell [1982], Hardouvelis [1984] and Frankel and Hardouvelis [1985]). Since we find that U.S. interest rates are negatively related to the value of the dollar (as do these other studies), our results provide additional evidence of the money target view for the U.S.. Note, since the parameter on German interest rates indicates a positive relationship between i and e, the regression results point to the inflation view for Germany.

consistent expectations.

In order to draw valid inferences from the FIML estimations of our reduction analysis, the variables contained in regressions 5 and 9 must be co-integrated, since the systems approach of Phillips [1991] applies only to systems of cointegrated variables.[24] In table 3 we report the results of unit root tests for the OLS residuals generated by the reduced models for the first two exchange rate regimes (RM1 and RM2) and the OLS residuals obtained from estimating the models of table 1 over the entire sample.[25] The results show that the unit root hypothesis is rejected at the 1 percent level for the residuals of RM1 and at the 5 percent level for the residuals of RM2; whereas, the unit root hypothesis cannot be rejected for the OLS residuals of any of the models in table 1 when estimated over the entire sample. Hence, the variables of both reduced models are found to be cointegrated within the regimes of relative parameter stability, thereby indicating that macroeconomic fundamentals do provide a long-run anchor (albeit a periodically shifting one) for exchange rate movements. Furthermore, the inability of earlier studies to find co-integrating vectors stems from the fact that they used sample periods covering the 1970s and 1980s in estimating their models.

The results in table 2 show that the co-integrating vectors in regressions 5 and 9 differ not only in terms of parameter values, but also in terms of the variables included in the long-run relationship, i.e., different variables matter during different time periods. This finding is suggestive that a primary factor behind the periodic shifting of the co-integrating vector is recurring shifts in expectations functions. It is clear that structural change can arise in a linear setup through one of

[24]Note, no inferences can be drawn from the standard errors of the OLS regressions due to the problem of spurious regressions.

[25]The co-integration results are unchanged if RM1 and RM2 are estimated using FIML.

two channels, either there is a change in the semi-reduced-form structure characterizing the economy (the semi-reduced-form channel) or there is a change in the way market agents form exchange rate expectations (the expectations channel). Although the findings in table 2 are consistent with the view that both channels may be important for the instability findings, they point to the expectations channel as the more important of the two. This is partly because the degree of instability on the part of the semi-reduced form would have to be extreme in order to explain the finding that different variables matter during different time periods. For example, such an explanation would require instability involving a different set of money demand determinants over time or a long-run real exchange rate that is constant during one period and then a function of cumulative trade balances in another. There is nothing in the literature indicating the plausibility of such changes in the underlying structure of the economy occurring on a regular basis. Furthermore, the finding that different variables matter during different time periods is consistent with the views developed in Froot, Scharfstein and Stein [1991] and Goldberg and Frydman [1993a,b] that market agents focus on different sets of fundamental variables during different time periods when forecasting exchange rate movements.[26]

There are several additional noteworthy aspects of the results in table 2. First, the problem of multicollinearity is evident during the first two exchange rate regimes. Some of the parameter estimates which are initially either insignificant or significant and of the wrong sign become significant and of the correct sign when insignificant variables are dropped from the regressions (e.g., German trade balances and U.S. income during the first and second exchange rate regimes

[26]In Goldberg and Frydman [1991] we show how shifts in expectations functions according to the theories consistent expectations framework can cause explanatory variables to come in and out of the reduced form. Note, that the Lucas critique cannot explain this finding, since with RE, all of the explanatory variables that are contained in the semi-reduced-form setup must enter the aggregate expectations functions. See footnote 18.

respectively). Although we have attempted to minimize the effects of multicollinearity, the presence of this problem suggests that the results in table 2 must be interpreted with care.

Second, we would expect to see that the FIML parameter estimates correspond well with those obtained using OLS, since the latter are super-consistent. Although this is mostly the case, the results in table 2 indicate that the OLS estimates involve considerable small sample bias in some cases (e.g. the parameter estimate on German cumulative trade balances (k) in regression (3)). As such, our results indicate that when working with I(1) variables, it is important to use estimation techniques that specify the unit roots explicitly.

Finally, note that when the D+ model is estimated over the entire sample the DW statistic is low, which is characteristic of a regression model involving I(1) variables which are not cointegrated (this is also the case with the other models of table 1). This is consistent with the results of previous studies (e.g., Frankel [1983], among many others). But, when the unstable nature of the co-integrating vector is taken into account and the D+ model is estimated within the regimes of relative parameter constancy the value of the DW statistic rises substantially in every exchange rate regime, which is indicative of stationary residuals, i.e., of cointegration. The D+ model for the first exchange rate regime experiences the greatest improvement, with a DW statistic of 1.86. Not only are the residuals from RM1 not characterized by the type of strong autocorrelation found with I(1) variables, but when they are tested for autocorrelation (by regressing them on ten lags) no lag is significant at the 10 percent level when based on either the t-test or F-test! These results indicate that the serially correlated residuals of regression 1 stem from a specification error and not from autoregressive errors in the true data generating process. As such, the common practice in the literature of correcting autocorrelated residuals using either Cochrane-Orcutt or Fair's instrumental variables technique when using the entire sample period is inappropriate (e.g., Frankel [1983] and

Boothe and Glassman [1987], among many others).

4.0 Out-of-Sample Fit

The findings of Meese and Rogoff [1983], which show the out-of-sample fit of popular structural models to be no better than that of the random walk model, are perhaps the findings most often cited in the literature as illustrating the failure of empirical exchange rate models. We show in this section, however, that these findings are, in large measure, the result of allowing the analysis to run past the end of one exchange rate regime and into the next. We find that if the "forecasting" experiment takes into account the shifting nature of the co-integrating vector, then the structural models outperform the random walk model by considerable margins.

In the Meese and Rogoff study structural models are first estimated over an initialization period that begins in March 1973 and ends in November 1976. These estimated models, along with the actual future values of the explanatory variables, are then used to generate forecasts at a number of horizons. Repeated forecasts are generated using rolling regressions and the out-of-sample analysis ends in June 1981. Hence, not only are the results of Meese and Rogoff [1983] based on an initialization period that includes the first transition period, but the forecasting experiment is allowed to run past the first exchange rate regime and into the second (i.e., it runs past two break points).

It is clear that in such a case, any measure of out-of-sample accuracy that involves averaging (such as root mean square error (RMSE)) will produce a time path that is at first relatively flat within the initial period of parameter constancy and then becomes positively sloped once the first break point is reached. In figure 2 we plot the RMSE produced by the HM and RM1 models at the 3-month forecasting horizon, where the initialization and forecasting periods are identical to those in

Meese and Rogoff [1983]. As figure 2 reveals, the out-of-sample accuracy of both the HM and RM1 models are rather consistent within the first regime of parameter constancy, but once the first break point is reached in October 1978, their ability to forecast completely breaks down.[27] If forecasting accuracy is measured as the average performance over the 1976-1981 period using the RMSE value generated as of June 1981, then not only will the performance of the structural models appear poor, but their forecasting ability prior to the first break point will be completely masked.

Summary statistics for the forecasting performance of the HM, RM1 and random walk models bear this point out. They are reported in table 4. The first half of table 4 replicates the dismal results of Meese and Rogoff [1983], where forecasting accuracy is measured as the average performance (in terms of RMSE) over the period from December 1976 to June 1981.[28] The figures in this first half of the table reveal that when compared to the out-of-sample fit of the random walk model the fit of the structural models are inferior by considerable margins at every forecasting horizon when based on the RMSE measure.[29] But, when the forecasting experiment is not allowed

[27]Note, plots of the RMSE produced at other forecasting horizons are nearly identical to those in figure 2.

[28]Mean forecast errors and mean absolute forecasts errors are also calculated. These results are found to mirror the results based on RMSE and as such they are not reported. The D+ model is also tested and found to perform slightly better than the reduced model for the first exchange rate regime and slightly worse during the second exchange rate regime.

[29]The results in the first half of table 4 are more extreme than the results reported in Meese and Rogoff [1983], which show the HM model to perform slightly better than the RW model at the 1 month horizon and slightly worse at the 3, 6 and 12 month horizons. This is because unlike in Meese and Rogoff [1983], the residuals of the structural models in table 4 are not corrected for first-order serial correlation (in effect adding the lagged value of the exchange rate as an additional regressor). (The DW statistic for the original HM model over the period from March 1973 through June 1981 is .15.) As mentioned in the preceding section the structural change and reduction results show that such serially correlated residuals stem from a specification error and not from autoregressive errors in the true data generating process. As such, it is inappropriate to correct for serial correlation using either Cochrane-Orcutt or Fair's instrumental variables technique (both of which are employed in Meese and Rogoff [1983]).

to run beyond the first exchange rate regime a very different story emerges. These results are presented in the second half of table 4. The figures on the top of each row are based on the same initialization period as that used in the first half of the table (the Meese and Rogoff [1983] case); whereas, the bottom figures are based on an initialization period that begins in July 1974 and ends in November 1976, thereby corresponding with the time period of the first exchange rate regime. In both cases forecasting begins in December 1976 and ends in September 1978 (as such the figures for the random walk model are the same in both cases). The figures in this second half of the table reveal that both the HM and the RM1 models outperform the random walk by substantial margins when based on RMSE at every forecasting horizon save the 1 month horizon. Furthermore, both models perform best at the longer forecasting horizons, suggesting that fundamentals play a greater role over the longer-run. For example, the RMSE of the RM1 model (4.67 percent) is almost one third that of the random walk model (13.59 percent). Hence, the large forecasting errors reported in Meese and Rogoff [1983] are largely the result of forecasting errors amassed after the end of the first exchange rate regime in September 1978.

Table 4 also reports on an alternative measure of out-of-sample accuracy not found in Meese and Rogoff [1983]. This alternative measure records the percentage of times that a model correctly predicts the direction of change of the exchange rate, which we call the right side of the market (RSM) measure. The RSM measure is useful not only because it lends itself easily to formal tests of significance when compared to a random process (via the binomial distribution), but also because it is consistent with a weaker form of rationality developed in Goldberg and Frydman [1993b]. Using the RSM measure, the same story emerges. When the forecasting experiment is allowed to run until June 1981, the ability of the structural models to predict the right side of the market is statistically indistinguishable from that of flipping a coin. But, when the out-of-sample analysis ends prior to the shift of the co-integrating vector in October 1978, the performance of the structural models improve markedly, especially at the longer forecasting horizons. Both the HM and the RM1 models are able to correctly predict the direction of change of the exchange rate 100 percent of the time at the 6, 9 and 12 month forecasting horizons.

In table 5 we examine the out-of-sample fit of the HM, RM2 and RW models during the second exchange rate regime, where we use an initialization period that runs from October 1979 through June 1982. Although not quite as impressive as the results for the first exchange rate regime, the results here lead to a similar conclusion. When the forecasting experiment is allowed to run until the end of the sample in March 1988, the performance of the structural models is everywhere inferior to that of the random walk. These results are presented in the first half of table 5. But, when the break in September 1984 is incorporated into the analysis, the performance of the structural models improves considerably, especially at the longer forecasting horizons. These results are presented in the second half of table 5. The figures on the bottom of each row are based on estimating an equation that includes among its regressors first differences of the dependent variable and each explanatory variable. The addition of such I(0) variables explains the autocorrelated residuals produced by the structural models during the second exchange rate regime and thus reflects the short run dynamics of the system. Note, that the RM2 model performs well relative to the random walk when based on both the RMSE and RSM measures, whereas the HM model performs well only with respect to its ability to predict the right side of the market.

5.0 Conclusion

We have shown that the failure of empirical exchange rate models is due in large measure to periodic shifts in the long-run relationship governing the exchange rate and macroeconomic fundamentals, i.e., to recurrent shifts in the co-integrating vector. Using structural change tests that search recursively for break points, rather than using tests that impose break points a priori, we found that the co-integrating vector implied by a composite monetary model experienced parameter shifts on five occasions over a sample period that included the 1970s and 1980s. These shifts in turn allowed us to distinguish three periods or regimes of relative parameter constancy along with three short-lived periods of transition preceding each regime. Within the regimes of relative parameter constancy we found that many of the fundamental variables of the monetary models were significant and entered with correct parameter signs. Furthermore, we found that the sets of significant variables in each regime were cointegrated, which implied the existence of stable long-run relationships operating in the foreign exchange market. Finally, all of the structural models examined outperformed the random walk model in out-of-sample fit by considerable margins within the separate regimes of stability, especially at the longer forecasting horizons. In some cases the structural models outperformed the random walk model by a margin of 70 percent in root mean square error and were able to predict the direction of change of the exchange rate 100 percent of the time at the 6, 9 and 12 month forecasting horizons. We found that the large forecasting errors reported in Meese and Rogoff [1983] were the result of allowing the forecasting experiment to run past the end of one exchange rate regime and into the next.

One of the questions we attempted to address in our study concerned the causes behind the periodic shifts in the co-integrating vector. Our findings suggested that shifts in expectations functions are an important factor behind the observed instability. First, we found that different sets of fundamental variables were significant during different time periods and second, we found that many of the observed points of parameter instability matched up well with time periods involving major changes in government policy. Our findings, therefore, lent support to the alternative approaches of Froot, Scharfstein and Stein [1992] and Goldberg and Frydman [1993a,b].

Another important question raised in our study concerned the degree to which our findings supported any of the monetary models of the exchange rate. In Goldberg and Frydman [1991] we attempt a detailed answer of this question, where we find that a sticky-price monetary model with theories consistent expectations is consistent with our empirical findings. The difficulty of reconciling the sticky-price monetary model with the data stems from our finding that different sets of fundamental variables are significant during different time periods. The theories consistent expectations framework not only allows for such behavior, but it implies that such behavior should occur through switching expectations functions.

Appendix

Description of Data

All data are monthly. The data set begins in March 1973 and ends in March 1988. The trade balance data are from the O.E.C.D. (M.E.I.) data bank. All other time series are from the IFS data bank.

e	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^*	M_1, end of month, in billions of local currency
k and k^*	in billions of local currency, cumulative sum started in January 1970

References

Alexander, D. and L. R. Thomas, III (1987), "Monetary/Asset Models of Exchange Rate Determination: How Well Have They Performed in the 1980s," *International Journal of Forecasting*, Vol. 3, pp. 53-64.

Bilson, J. F. O. (1978), "Recent Developments in Monetary Models of Exchange Rate Determination," *IMF Staff Papers*, June, pp. 201-223.

Bilson, J. F. O. (1978), "The Deutsche Mark/Dollar Rate — A Monetary Analysis," in K. Brunner and A. H. Meltzer (eds.), *Policies for Employment, Prices and Exchange Rates*, Carnegie-Rochester Conference 11, Amsterdam: North-Holland.

Boothe, P. and D. Glassman (1987), "Off the Mark: Lessons for Exchange Rate Modelling," *Oxford Economic Papers*, 39, pp. 443-457.

Boughton, J. M. (1987), "Tests of the Performance of Reduced-Form Exchange Rate Models," *Journal of International Economics*, 23, pp. 41-56.

Brown, R. L., J. Durbin and J. M. Evans (1975), "Techniques for Testing the Constancy of Regression Relationships Over Time (with discussion)," *Journal of the Royal Statistical Society*, B, 37, pp. 149-192.

Chinn, M.D. (1991), "Some Linear and Nonlinear Thoughts on Exchange Rates," *Journal of International Money and Finance*, 10, pp. 214-230.

Cornell, B. (1982), Money Supply Announcements, Interest Rates, and Foreign Exchange," *Journal of International Money and Finance*, August, pp. 201-208.

Dickey, D. A. and W. A. Fuller (1981), "Likelihood Ratio Statistics for Autoregressive Time Series With a Unit Root," *Econometrica*, 49, pp. 1057-1072.

Dornbusch, R. (1980), "Exchange Rate Economics: Where Do We Stand?," *Brookings Papers on Economic Activity*, 1, pp. 143-94.

Dornbusch, R. (1976), "Expectations and Exchange Rate Dynamics," *Journal of Political Economy*, December, pp. 11161-1174.

Engel, C. M. and J. D. Hamilton (1990), "Long Swings in the Exchange Rate: Are They in the Data and Do Markets Know It?," *American Economic Review*, 80, September, pp. 689-713.

Engel, C. M. and J. A. Frankel (1984), "Why Interest Rates React to Money Supply Announcements: An Explanation from the Foreign Exchange Market," *Journal of Monetary Economics*, 13, January, pp.31-39.

Engel, C. M. and J. A. Frankel (1982), "Why Money Announcements Move Interest Rates; An Answer from the Foreign Exchange Market," NBER Working Paper No. 1049, December.

Engel, R. F. and C. W. J. Granger (1987), "Co-Integration and Error Correction: Representation, Estimation, and Testing," *Econometrica*, 55, 2, March, pp. 251-276.

Evans, G. (1986), "A Test for Speculative Bubbles and the Sterling-Dollar Exchange Rate: 1981-84," *American Economic Review*, 76, pp. 621-636.

Frankel, J. A. (1983), "Monetary and Portfolio Balance Models of Exchange Rate Determination," in J. Bhandari and B. Putnam (eds.), *Economic Interdependence and Flexible Exchange Rates*, Cambridge: MIT press.

Frankel, J. A. (1982), "The Mystery of the Multiplying Marks: A Modification of the Monetary Model," *The Review of Economics and Statistics*, August, pp. 515-519.

Frankel, J. A. (1979), "On the Mark: A Theory of Floating Exchange Rates Based On Real Interest Differentials," *American Economic Review*, September, pp. 610-622.

Frankel J. A. and K. A. Froot (1987), "Explaining the Demand for Dollars: International Rates of Return, and the Expectations of Chartists and Fundamentalists," Economics Working paper No. 8603, U.C. Berkeley. In R. Chambers and P. Paarlberg (eds.), *Agriculture, Macroeconomics and the Exchange Rate, Boulder*, Colorado: Westview Press.

Frankel J. A. and G. A. Hardouvelis (1985), "Commodity Prices, Money Surprises and Credibility," *Journal of Money, Credit, and Banking, November*, pp. 425-438.

Frenkel, J. A. (1976), "A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence," *Scandinavian Journal of Economics*, 78, pp. 200-224.

Frenkel, J. A. and C. A. Rodriguez (1982), "Exchange Rate Dynamics and the Overshooting Hypothesis," *International Monetary Fund Staff Papers*, 29, pp. 1-30.

Froot, K.A., D.S. Scharfstein and J.C. Stein (1992), "Herd on the Street: Informational Inefficiencies in a Market with Short-Term Speculation," *Journal of Finance*, September, pp. 1461-1484.

Frydman, R. (1983), "Subjective Forecast and the Output-Inflation Relationship," C.V. Starr Center for Applied Economics Research Report.

Fuller, W. A. (1976), *Introduction to Statistical Time Series*, New York: Wiley.

Goldberg, M.D. (1993a), "Long-Run Non-Neutrality in a Standard Sticky-Price Monetary Model of the Exchange Rate," C.V. Starr Center for Applied Economics Working Paper.

Goldberg, M. D. (1993b), "Distinct Reduced-Form Parameters in a Standard Sticky-Price Monetary Model of the Exchange Rate," C.V. Starr Center for Applied Economics Working Paper.

Goldberg, M. D. (1991), "Reconsidering the Basic Relationships Between Exchange Rates, Exchange Rate Expectations and Macroeconomic Fundamentals," PH.D. Dissertation, New York University, October.

Goldberg, M.D. and R. Frydman (1991), "Re-Examining the Empirical Performance of the Monetary Models of the Exchange Rate: A Problem of Structural Change," C.V. Starr Center for Applied Economics Working Paper #91-69, December.

Goldberg, M. D. and R. Frydman (1993a), "Theories Consistent Expectations and Exchange Rate Dynamics," forthcoming in H. Frisch and A. Wörgöter (eds.), *Open Economy Macroeconomics*, MacMillan, 1993.

Goldberg, M. D. and R. Frydman (1993b), "Qualitative Rationality and Behavior in the Foreign Exchange Market," C.V. Starr Center for Applied Economics Working Paper.

Granger, C. W. J. (1986), "Developments in the Study of Cointegrated Economic Variables," *Oxford Bulletin of Economics and Statistics*, 48, 3, pp. 213-228.

Granger, C. W. J. (1983), "Co-Integrated Variables and Error-Correcting Models," UCSD Discussion Paper 83-13a.

Granger, C.W. J. and P. Newbold (1974), "Spurious Regressions in Econometrics," *Journal of Econometrics*, Vol. 2, pp. 111-120.

Hardouvelis, G. A. (1984), Market Perceptions of the Federal Reserve Policy and the Weekly Monetary Announcements," *Journal of Monetary Economics*, September, pp. 225-240.

Hendry, D. F. (1986), "Econometric Modelling with Cointegrated Variables: An Overview," *Oxford Bulletin of Economics and Statistics*, 48, 3, pp. 201-212.

Hooper , P. and J. Morton (1982), "Fluctuations in the Dollar: A Model of Nominal and Real Exchange Rate Determination," *Journal of International Money and Finance*, April, pp. 39-56.

Johansen, S. (1988), "Statistical Analysis of Cointegration Vectors," *Journal of Economic Dynamics and Control*, 12, pp. 231-254.

Kaminsky, G. (1993), "Is There a Peso Problem?: Evidence From the Dollar/Pound Exchange Rate, 1976-1987," *American Economic Review*, June, pp. 450-472.

Kirman, A. P. (1990), "On Ants and Markets," mimeo European University Institute, January.

Meese, R. (1986), "Testing for Bubbles in Exchange Markets: A Case of Sparkling Rates?," *Journal of Political Economy*, vol. 94, no. 2, pp. 345-373.

Meese, R. and K. Rogoff (1988), "Was It Real? The Exchange Rate-Interest Differential Relation Over the Modern Floating-Rate Period," *Journal of Finance*, September, pp. 933-948.

Meese, R. and K. Rogoff (1983), "Empirical Exchange Rate Models of the Seventies: Do They Fit Out of Sample?," *Journal of International Economics*, February, pp. 3-24.

Meese, R. and A. Rose (1991), "An Empirical Assessment of Non-Linearities in Models of Exchange Rate Determination," *Review of Economic Studies*, 58, pp. 603-619.

Nicoletti, G. (1989), "Saving and Consumption in High Deficit Countries: Discriminating Between Contingent and Permanent Effects of Government Financing Policies in Italy and Belgium," chapter III Ph.d. Dissertation.

Papell, D. H. (1988), "Expectations and Exchange Rate Dynamics After a Decade of Floating Rates," *Journal of International Economics*, Vol. 25, pp. 303-317.

Perron, P. (1988), "Trends and Random Walks in Macroeconomic Time Series," *Journal of Economic Dynamics and Control*, 12, pp. 297-332.

Phillips, P. C. B. (1991), "Optimal Inference in Cointegrated Systems," Cowles Foundation Working Paper No. 42, February.

Phillips, P. C. B. (1987), "Time Series Regression with Unit Roots," *Econometrica*, 55, pp. 277-302.

Phillips, P. C. B. (1986), "Understanding Spurious Regressions in Econometrics," *Journal of Econometrics*, Vol. 33, pp. 311-340.

Phillips, P. C. B. and B. E. Hansen (1988), "Statistical Inference in Instrumental Variables regression with I(1) Processes," Cowles Foundation Working Paper No. 43, February.

Phillips, P. C. B. and S. Ouliaris (1988), "Asymptotic Properties of Residual Based Tests for Cointegration," Cowles Foundation Working Paper No. 847-R, July.

Schinasi, G. J. and P. A. V. B. Swamy (1989), "The Out-of-Sample Forecasting Performance of Exchange Rate Models When Coefficients are Allowed to Change," *Journal of International Money and Finance*, September, pp. 375-390.

Stock, J. H. (1987), "Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors," *Econometrica*, 55, pp. 1035-1056.

Table 1
The Predictions of the Asset Market Models[a]

Models[b]

Variables[c]	FB	DF	HM	D+
m				+
m^*				-
$m - m^*$	+	+	+	
y				-
y^*				+
$y - y^*$	-	-	-	
i				+/-
i^*				+/-
$i - i^*$	+	-	-	
π				+
π^*				-
$\pi - \pi^*$		+	+	
k			-	-
k^*			+	+
constant	+/-	+/-	+/-	+/-

Notes:

a. Symbols indicate parameter sign predictions of the structural models listed. The symbol "+" denotes the prediction of a positive parameter value and the symbol "-" denotes the prediction of a negative value. Entries that include both symbols indicate that the model's prediction is ambiguous for that parameter value.

b.Model definitions: FB denotes the Frenkel-Bilson model, DF denotes the Dornbusch-Frankel model, HM denotes the Hooper-Morton model and D+ denotes the Dornbusch et al. model.

c. Variable Definitions: m, p, y are the log levels of domestic (German) money supply, goods prices and real income respectively, i is the level of the domestic short-term interest rate, π is the expected long-run domestic inflation rate, k is the de-trended level of domestic cumulative trade balances and a "*" denotes foreign-country (U.S.) variable.

Table 2
Regression Analysis
FIML and OLS[a]

Explanatory[b]	Lagged values[c]	Regressions									
		1973:3-1988:3	1974:7 - 1978:9				1979:10 - 1984:8				1985:9-1988:3
		OLS	OLS		FIML		OLS		FIML		OLS
		1	2	3	4	5	6	7	8	9	10
m	2,6	-.777	-.007		-.045		-.103	-.467	-.117	-.407	.092
	2,4	.214	.123		.086		.174	.182	.112	.121	.299
m*	1,2,6	1.832	.049		-.023		.137	.692	.283	.912	-.383
	1	.210	.233		.165		.288	.204	.188	.144	.512
y	1	-.931	-.446	-.732[d]	-.970[d]	-.916[d]	-.130	-.622[d]	.319[d]	-.873[d]	.452
	1	.297	.186	.142	.130	.010	.226	.154	.145	.105	.530
y*	1,3	1.017	1.052	.930[d]	1.115[d]	1.115[d]	-.141	.439[d]	-.415[e]	.310[d]	-2.093
	1	.224	.201	.157	.141	.110	.225	.168	.163	.117	1.463
i	1	-3.558	-2.105	-1.233[d]	-1.169[d]	-1.325[d]	2.008	1.568[d]	2.566[d]	1.803[d]	2.062
	1,2,3	.592	.659	.326	.465	.230	.573	.396	.385	.266	.374
i*	3,8	.751	2.778	3.283[d]	2.685[d]	2.232[d]	1.123	.702[d]	1.312[d]	.701[d]	3.071
	2,4,6	.575	.653	.555	.462	.379	.312	.341	.204	.228	.342
π	1	.791	.065		-.016		.095		.002		.633
	1	.123	.116		.082		.114		.023		.380
π*	1	-.397	.055		.008		.284	-.375[d]	-.179[d]	-.285[d]	-.289
	1	.048	.058		.041		.074	.061	.056	.042	.274
k	2,6	-.287	-.050	-.003[d]	-.019	-.067[d]	.101		.143[e]		.078
	3,5	.058	.088	.034	.063	.024	.108		.071		.396
k*	1,5	-.024	.693	.706[d]	.670[d]	.618[d]	-.078		-.069[g]		.079
	3,5	.026	.080	.050	.056	.036	.058		.038		.260
constant		-5.348	-1.972		.667		1.110		2.183[f]		10.363
		1.476	1.199		.856		1.769		1.137		7.855
R^2		.773	.972	1.000			.967	.998			.966
DW		.51	1.85	1.80			1.45	1.33			1.64

Notes:

a. Estimations are based on either OLS or on the full information maximum likelihood (FIML) approach of Phillips [1991]. (See section 3.1.) The dependent variable is the German mark-U.S. dollar exchange rate. No lagged values of u_t are added to the cointegrating equation in regressions (4) and (5) because u_t is found to be i.i.d. for the first exchange rate regime. One lagged value of u_t is added to the cointegrating equation in regressions (8) and (9) in order to explain u_t up to an i.i.d. error process for the second exchange rate regime (see equation system (2)).

b. See Table 1 for variable definitions. All data are monthly.

c. Figures on the top of each row indicate the lagged values of Δx_t required in order to explain v_t up and to an i.i.d. error process for the first exchange rate regime (see equation system (2)). Figures on the bottom are for the second exchange rate regime.

d. Significant at the 95 percent level and of a correct sign.

e. Significant at the 95 percent level and of an incorrect sign.

f. Significant at the 90 percent level and of a correct sign.

g. Significant at the 90 percent level and of an incorrect sign.

Table 3
Residual Based Tests of Cointegration
ADF and Z Tests

Test Statistics[a]

Models[b]	p[c]	n[d]	$t_{\hat{\alpha}}$	$z(t_{\hat{\alpha}})$	$z(\hat{\alpha})$
RS1	3	6	-2.08	-6.00[e]	-41.97[f]
RS2	3	7	-3.73	-5.52[e]	-39.58
FB	5	3	-1.00	-1.59	-6.34
DF	3	4	-2.15	-1.93	-9.36
HM	5	6	-3.22	-4.57	-37.13
D+	7	10	-3.29	-5.19	-46.27

Notes:

a. All co-integrating vectors are estimated using OLS. (When FIML is used to estimate the co-integrating vectors of RM1 and RM2 similar results are obtained.) The test statistics t_{α} and $Z(t_{\alpha})$ are the augmented Dickey and Fuller (ADF) and Phillips and Perron Z t-statistics for the size of the root respectively, whereas the test statistic $Z(\alpha)$ is the Phillips and Perron Z α-statistic for the size of the root (see Fuller [1976] and Dickey and Fuller [1979] and Perron [1987]). Critical values depend on the number of regressors (n) in the cointegrating relationship. Critical values are obtained from Phillips and Ouliaris [1988] for the cases where n is less than or equal to 5. For cases where n is greater than 5 critical values are obtained through extrapolation. Critical values also depend on whether the co-integrating vector includes a constant, which is the case for the FB, DF, HM and D+ models.

b. Definitions: RM1 and RM2 denote the reduced models for the first and second exchange rate regimes respectively (regressions (3) and (7) in table 2). FB is the Frenkel-Bilson model, DF is the Dornbusch-Frankel model, HM is the Hooper and Morton model and D+ is the Dornbusch et. al. model (see table 1 for more detail on these models). The co-integrating vectors implied by these last four models are estimated over the entire sample period.

c. Figures indicate the number of lagged first difference terms found to be significant in obtaining the ADF statistics, i.e., ($\Delta e_t = \alpha e_{t-1} + \Sigma \Delta e_{t-i} + \epsilon_t$), where i = 1, p).

d. Figures indicate the number of regressors in the cointegrating vector.

e. Significant at the 95 percent confidence level.

f. Significant at the 90 percent confidence level.

Table 4
Out-of-Sample Fit: First Exchange Rate Regime
Root Mean Square Forecast Errors (RMSE) and Right Side of the Market (RSM)[a]

	The Meese and Rogoff Case[b]						The First Exchange Rate Regime[c]					
Forecasting Horizon (in months)	RMSE			RSM			RMSE			RSM		
	RW	HM	RM1	HM	RM1	Nob[d]	RW	HM	RM1	HM	RM1	Nob[d]
1	3.63	7.32	10.83	51	47	55	2.15	2.64	2.54	64	59	22
								3.08	2.28	59	59	
3	5.83	9.65	13.30	58	58	53	4.32	3.24	3.20	80[e]	80[e]	20
								3.68	2.87	75[e]	80[e]	
6	8.93	12.69	16.36	68	68	50	6.77	3.46	3.27	100[e]	100[e]	17
								3.65	2.47	100[e]	100[e]	
9	11.08	15.48	19.48	70	70	47	10.14	4.37	3.95	100[e]	100[e]	14
								4.03	3.56	100[e]	100[e]	
12	13.08	18.63	23.59	72	72	44	13.59	4.92	5.36	100[e]	100[e]	11
								6.82	4.67	100[e]	100[e]	

Notes:

a. Since all models are for the logarithm of the exchange rate, the RMSE figures are roughly in percentage terms. RSM figures denote the percentage of times models are able to correctly predict the direction of change of the exchange rate. The two structural models - the original Hooper and Morton Model (HM) and the reduced model for the first exchange rate regime (RM1) - are estimated using OLS. RW denotes the random walk model.

b. Figures are based on an initialization period that begins in March 1973 and runs through November 1976. Forecasting begins in December 1976 and runs until June 1981.

c. The figures on the top of each row are based on an initialization period that begins in March 1973 and runs through November 1976, whereas the figures on the bottom are based on an initialization period that begins in July 1974 and runs through November 1976. In both cases forecasting begins in December 1976 and ends in September 1978.

d. Number of forecasting observations.

e. Significant at the 5 percent level when based on the binomial distribution.

Table 5
Out-of-Sample Fit: Second Exchange Rate Regime
Root Mean Square Forecast Error (RMSE) and Right Side of the Market (RSM)[a]

Forecasting Horizon (in months)	End of Forecasting Period: March 1988						The Second Exchange Rate Regime[b]					
	RMSE			RSM			RMSE			RSM		
	RW	RM2	HM	RM2	HM	Nob[c]	RW	RM2	HM	RM2	HM	Nob[c]
1	3.51	11.88	10.10	35	42	69	2.80	3.95	4.30	46	54	26
								3.27	4.45	65[d]	58	
3	6.46	13.85	12.96	51	40	67	4.57	4.39	5.92	79[d]	58	24
								3.60	6.18	79[d]	54	
6	10.62	16.60	12.96	53	44	64	5.69	4.54	7.44	76[d]	67	21
								3.55	8.27	86[d]	71[d]	
9	14.56	19.38	22.15	48	51	61	6.90	5.17	8.63	61	94[d]	18
								4.12	10.38	67	94[d]	
12	18.59	22.48	27.26	52	44	58	8.41	5.89	11.12	87[d]	93[d]	15
								4.98	13.98	87[d]	100[d]	

Notes:

a. Since all models are for the logarithm of the exchange rate, the RMSE figures are roughly in percentage terms. RSM figures denote the percentage of times models are able to correctly predict the direction of change of the exchange rate. All figures are based on an initialization period that begins in November 1979 and runs through June 1982. Forecasting begins in July 1982. The two structural models - the original Hooper and Morton Model (HM) and the reduced model for the second exchange rate regime (RM2) - are estimated using OLS. RW denotes the random walk model.

b. Figures on the top of each row are for the models indicated, whereas the figures on the bottom of each row are based on adding to the indicated model first differences of all the dependent variable and all of the explanatory variables. The forecasting period ends in August 1984.

c. Number of forecasting observations.

d. Significant at the 5 percent level when based on the binomial distribution.

Figure 1
Exchange Rate Regimes and Transition Periods
Sample Period 1973:03 - 1988:03

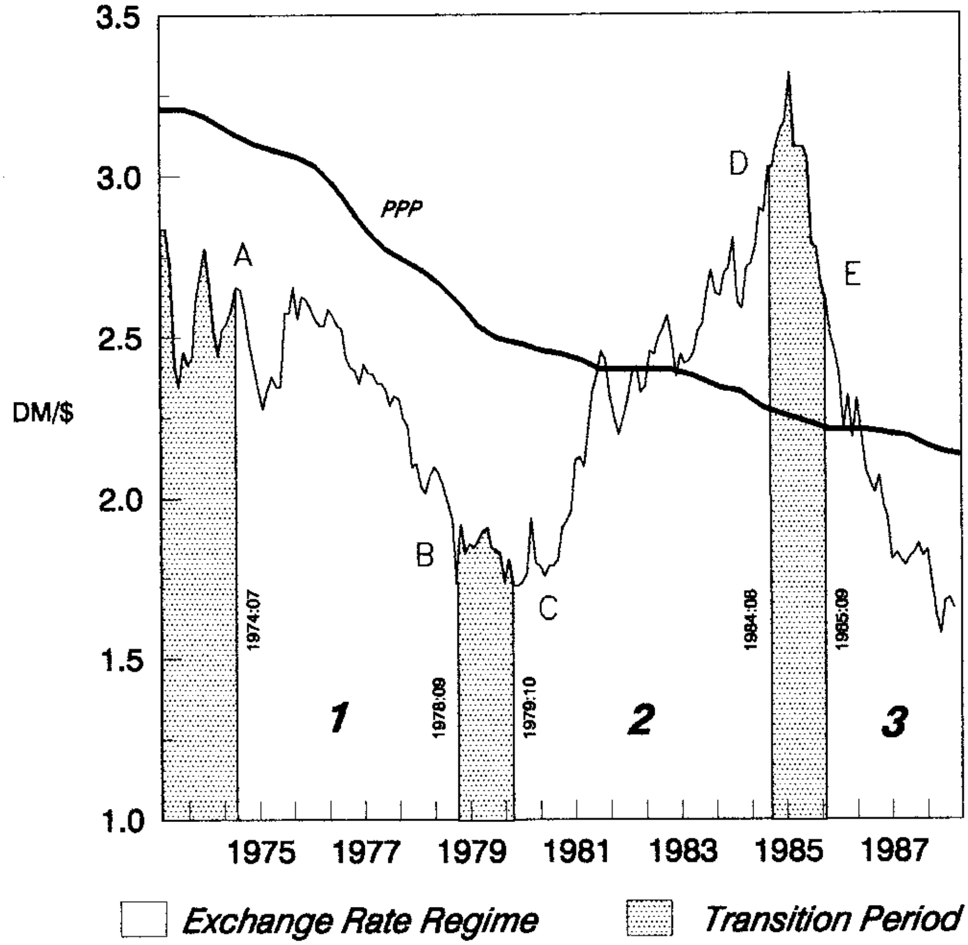

Figure 2

The Meese and Rogoff Case