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Predicting real exchange rates from real interest rate differentials and net foreign asset stocks: evidence for the mark/dollar parity

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**Predicting Real Exchange Rates from
Real Interest Rate Differentials and
Net Foreign Asset Stocks:
Evidence for the Mark/Dollar Parity**

by

Carsten-Patrick Meier

December 1999

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Abstract:

When nontraded goods prices are accounted for consistently and genuine stock data on bilateral foreign asset holdings is employed, a modified sticky-price exchange rate model by far outperforms the benchmark random walk-model in empirically forecasting the D-mark/dollar parity out of sample. Superior forecast performance holds both over long horizons and from the first step. Extending the sample back to the Bretton Woods period leaves the model's parameters and its performance virtually unaffected. By implication, the explanatory variables of the model show a pattern of exchange rate regime-dependent volatility that is similar to that of the real exchange rate itself.

Keywords:

Real exchange rates; Real interest rates; Net foreign assets; Non-tradables prices; Fixed/floating exchange rate regimes

JEL classification: F31, F32

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When nontraded goods prices are accounted for consistently and genuine stock data on bilateral foreign asset holdings is employed, a modified sticky-price exchange rate model by far outperforms the benchmark random walk-model in empirically forecasting the D-mark/dollar parity out of sample. Superior forecast performance holds both over long horizons and from the first step. Extending the sample back to the Bretton-Woods-period leaves the model's parameters and its performance virtually unaffected. By implication, the explanatory variables of the model show a pattern of exchange rate regime-dependent volatility that is similar to that of the real exchange rate itself.

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1 Introduction

This paper investigates the behavior of the quarterly real exchange rate of the U.S. dollar vis-à-vis the D-mark over the recent floating period. I find that when the prices of non-traded goods are accounted for consistently and genuine stock data on bilateral net foreign asset holdings is employed, a reduced-form time series equation based on the sticky-price monetary model proposed of Dornbusch (1976) and Frankel (1979) by far outperforms the benchmark random walk-model in out-of-sample forecasting. Over longer forecast horizons, the model's root mean square forecast error is only a quarter of that of the naive forecast; still, superior forecast performance is not confined to long horizons but holds from the first forecast step. Rival specifications, some of which are proposed in the literature, can be shown to be dominated in terms of both forecast quality and in-sample fit. With coefficient estimates stable and in line with theoretical prejudices the model is capable of explaining roughly sixty percent of the total quarterly variation of the dollar/mark real exchange rate over the floating rates period. Under reasonable assump-

tions, the results support the hypothesis that the nominal exchange rate overshoots its long-run level in the short run, but they also imply that real exchange rate adjustment typically proceeds more gradually than standard sticky-price theories would predict (Eichenbaum and Evans 1995). When the sampling period is extended to include ten years of observations from the Bretton-Woods system of fixed exchange rates, I find the parameter estimates virtually unaffected; the model's in-sample fit and its out-of-sample forecast performance even improve. By implication, the fundamentals used to explain the real exchange rate show a pattern of exchange rate regime-dependent volatility that is similar to that of the real exchange rate itself. So, in contrast to most published work, the model has the desirable property (see Flood and Rose 1995) of accounting for the stylized fact that the volatility of the real exchange rate is substantially higher under a regime of flexible nominal exchange rates than under a fixed rate regime.

Starting point of my analysis is the traditional sticky-price model of exchange rate behavior incorporated in the works of Mundell (1961) and Fleming (1962) and developed in detail by Dornbusch (1976) and Frankel (1979). Distinctive features of the approach are on the one hand its focus on short-run deviations of the real exchange rate from its long-run level which is defined to be the simple purchasing power parity (PPP) level and on the other the assumption of differential speeds of adjustment in markets for financial assets and in markets for goods which give rise to the characteristic dynamics of the real exchange rate. Financial markets are assumed to adjust instantaneously to disturbances, so uncovered interest parity (UIP) holds at every point in time. Goods prices, in contrast, adjust slowly to their long-run level, thus PPP only obtains in the long run. Movements of the real exchange rate under these assumptions occur as a result of international differences in real interest rates. Yet, since other factors to affect the real exchange rate are not considered such deviations from PPP are purely temporary phenomena caused by the short-run stickiness of goods prices.

Direct structural econometric support for the standard sticky-price model has been weak — as has been the econometric support for other exchange rate models based on fundamentals. Some promising recent attempts notwithstanding, the finding of Meese and Rogoff (1983, 1988) of a poor out-of-sample forecasting performance of structural exchange rate models in relation to a naive random walk has not yet been overturned.¹ Empirical support for the main assumptions of the standard sticky-price ap-

¹ Chinn and Meese (1995) and Mark and Choi (1997) show that the random walk model may be outperformed over long forecast horizons by augmented versions of the standard model. Mark (1995) as well as MacDonald and Taylor (1993, 1994) and MacDonald and Marsh (1997) shows that the flexible-price monetary model may have superior forecast power over long horizons. The gain in forecast accuracy is, however, rather limited. Moreover, there has been some criticism regarding the interpretation of long-horizon predictability tests see Kilian (1999).

proach, in contrast, is relatively strong. Mussa's (1986) finding of a high short-run correlation between nominal and real exchange rates, the large and persistent deviations from the law of one price documented by Engel (1993), Froot, Kim and Rogoff (1995) and Engel and Rogers (1996) and the results of some direct investigations into the subject by Cecchetti (1986), Kashyap (1995) and Blinder et al. (1998) together lend strong support for the assumption that goods prices are sticky in the short run. Moreover, purchasing power parity as a long-run proposition, another key assumption of the sticky-price monetary model, also seems to be well established empirically by now: A large number of papers document that, in general, goods prices and exchange rates tend to converge to PPP in the long-run (see Rogoff 1996 for a survey). Interestingly, however, while being overwhelming for most industrialized countries' currency pairs, the evidence in favor of long-run PPP seems to be less clear cut for U.S. dollar parities, at least for the recent floating rates period: a number of panel studies find that mean-reversion of the real exchange rate is accepted much more easily empirically when the D-mark, rather than the U.S. dollar, is used as the base currency (see e. g. Abuaf and Sweeney 1996, Papell and Theodoridis 1998, Lothian 1998, Canzoneri, Cumby and Diba 1999). Taken together, I read the empirical evidence for the standard sticky-price model as indicating that the model generally is a suitable framework for modeling industrialized countries' exchange rates. The assumption of a constant long-run real exchange rate, however, is a rather strong one which apparently it does not hold universally. Therefore, it seems worthwhile to make the standard model more flexible by allowing for permanent or long-run deviations from PPP.

For the United States, casual observation suggests that one source of long-run deviations from PPP may have been current account imbalances. Over the past twenty years, the United States have been accumulating foreign debt in an unprecedented magnitude while current account imbalances between other industrialized countries remained modest, fitting the evidence that for most other countries' currencies PPP seems to hold in the long-run. Theoretically, current account imbalances may affect the exchange rate either via their impact on expectations regarding the long-run sustainability of a given net foreign asset position or via their effect on portfolio risk considerations (Hooper and Morton 1982).² Both views suggest that under reasonable assumptions the volume of net foreign debt should cause the currency of the debtor country to depreciate relative to its PPP-value. Econometric support for this hypothesis has been rare, however, which may have to do with the lack of genuine stock data on bilateral net foreign asset holdings.³

² See also Kouri (1976), Dornbusch and Fischer (1980), Hooper and Morton (1982), Mussa (1984), Engel and Flood (1985), Krugman (1985, 1990), Meese and Rogoff (1988), and Faruq e (1995).

³ Masson, Kremers and Horne (1994) have compiled data on the net foreign asset positions of the United States, Japan and Germany, respectively, against the rest of the world. These multilateral stock data may

Whereas bilateral flows of goods and services on the one hand and capital (or claims) on the other are regularly reported in national balance of payments statistics, this is not the case for their accumulated stock. Cumulating the data on bilateral flows will only by chance fill the gap appropriately since it does not take into account valuation changes caused by changing prices of shares, bonds, real estates etc. and exchange rates, which can be quite important in practice.⁴ Estimating augmented versions of the standard sticky-price model for bilateral real exchange rates on the basis of these cumulated flow data, therefore, has not been very successful (Meese and Rogoff 1988, Edison and Pauls 1993, Chinn and Meese 1995). In the present paper, in contrast, I use genuine stock data on the bilateral net foreign asset foreign asset position between the United States and Germany, constructed from the German current account statistics, to estimate a sticky-price model for the D-mark/dollar parity. I find clear evidence for a negative impact of U.S. foreign debt on the real external value of the dollar.

A second source of changes in the long-run real exchange rate may be differences in the (relative) prices of non-traded goods. Under the conventional definition the real exchange rate measures the overall price level of one country relative to that of another country. International differences in the growth rates of nontraded goods prices, which by definition are not equalized by international arbitrage, are reflected in changes in the long-run real exchange rate.⁵ Empirically, roughly half of all goods and services produced today in a typical industrialized country are not traded internationally (DeGregorio, Giovannini and Wolf 1994). Moreover, a number of studies have indeed found growth rates of nontraded goods prices to be different across industrialized countries (Marston 1987, DeGregorio, Giovannini and Wolf 1994, Canzoneri, Cumby and Diba 1999), so there is scope for effects of nontraded goods prices on the real exchange rate. Direct evidence for these effects also comes from Engel (1999) for a number of U.S. dollar real exchange rates, who also finds that their magnitude is rather small relative to other effects.⁶ Attempts to account for relative nontraded goods prices or their determinants in structural econometric exchange rate models, however, again have shown only limited success (Throop 1993, Chinn and Meese 1995, Mark and Choi 1997). Yet, as I will argue below, the relative unsatisfactory results of previous studies are likely to be due to the inconsistent way in which nontraded goods prices have been integrated into the standard model. Using a slightly modified approach, I find that ac-

be used for studying effective exchange rates as done by Faruq e (1995) and MacDonald (1995); they are obviously inadequate when the focus is on bilateral parities.

⁴ For instance, for the years of 1995 and 1996 Deutsche Bundesbank (1998) estimates that only roughly half of the fall of Germany's net foreign asset position against the rest of world has to be attributed to the deficit in the current account in that years, the other half was caused by valuation effects.

⁵ See Froot and Rogoff (1995) and Rogoff (1996) for a survey on this literature.

⁶ The relevance of Engel's (1999) finding for the present study is discussed below.

counting for nontraded goods prices is in fact crucial for modeling the behavior of the real exchange rate of the D-mark vis-à-vis the U.S. dollar.

The rest of the paper is organized as follows. The modified sticky-price monetary model the empirical work is based on is derived in section 2. Section 3 lays out the data set and presents the results of some preliminary analysis. Section 4 clarifies the econometric approach and presents the results of estimating the model on data from the floating exchange rates-period together with some diagnostic checks. Section 5 assesses the model's relative out-of-sample forecast performance and compares its specification to alternative models, some of which have been used previously in the literature. Section 6 extends the model's database to include observations from the Bretton-Woods period, and section 7 concludes.

2 The Modified Sticky-Price Model

2.1 The Standard Approach

Meese and Rogoff (1988) represent the essence of the Dornbusch-Frankel model the following way: The natural logarithm of the real exchange rate q_t is defined as $q_t = e_t + p_t - p_t^*$ where e is the log of the nominal exchange rate (D-marks per U.S. dollars), p and p^* denote, respectively, U.S. and German aggregate price levels such as consumer prices or deflators for domestic aggregate output and the subscript t indicates the period of time. PPP implies that the real exchange rate is constant⁷, $q_t = \bar{q}$. Deviations of the real exchange rate from PPP are assumed to have the simple dynamics $E_t(q_{t+k} - \bar{q}) = \alpha_k(q_t - \bar{q})$, where E_t denotes expectations conditional on time- t -information and $0 < \alpha_k < 1$ is a speed-of-adjustment parameter that reflects the degree of price stickiness in goods markets. Solving for the current real exchange rate one finds

$$q_t = \bar{q} + \frac{1}{1 - \alpha_k} [E(q_{t+k} - q_t)]. \quad (1)$$

Uncovered interest parity implies ${}_k i_t - {}_k i_t^* = E_t(e_{t+k}) - e_t$, where ${}_k i_t$ is the U.S. and ${}_k i_t^*$ the German k -period nominal interest rate at time t . If PPP obtains in the long run, the expected long-run depreciation will be equal to the difference in the expected changes of price levels at home and abroad ${}_k i_t - {}_k i_t^* = E_t(p_{t+k}^* - p_t^*) - E_t(p_{t+k} - p_t)$ implying that k -period ex ante real interest rates ${}_k r_t = {}_k i_t - E_t(p_{t+k} - p_t)$ are equalized across the two countries in the long run. In the short run, however, sluggish price adjustment will cause the nomi-

⁷ The formulation assumes that transport costs as well as international differences in preferences for the various goods prevent price indices from being exactly equal, allowing for a non-zero logarithm of the real exchange rate.

nal interest rate differential to deviate from the expected inflation differential. As a result, the ex ante real interest rate differential is ${}_k r_t - {}_k r_t^* = {}_k i_t - {}_k i_t^* - E_t(p_{t+k} - p_t) + E_t(p_{t+k}^* - p_t^*) \neq 0$, which may be simplified to yield ${}_k r_t - {}_k r_t^* = E_t(q_{t+k} - q_t)$ by use of the UIP-relation and the definition of the real exchange rate $q_t = e_t + p_t^* - p_t$. Inserting this expression into (1) gives the characteristic relation of the sticky-price model between the real exchange rate and the real interest differential

$$q_t = \bar{q} + \frac{1}{1 - \alpha_k} ({}_k r_t - {}_k r_t^*) \quad (2)$$

which asserts that a transitory deviation of the current real exchange rate from its PPP level can be explained by the difference between the real rates of interest in the two countries involved.

2.2 Accounting for the Prices of Nontraded Goods

To model the impact of the prices of nontraded goods on the real exchange rate, let g be the share of these 'traded goods' in the home country's aggregate price level p_t , p_t^T their price aggregate and p_t^N the price aggregate of the non-traded goods, such that the U.S. aggregate price level is defined as $p_t = g p_t^T + (1-g) p_t^N$ and the German price level as $p_t^* = g^* p_t^{T*} + (1-g^*) p_t^{N*}$. Inserting these definitions into the definition of the real exchange rate, the (log) long-run real exchange rate based on aggregate price levels can then be expressed as the sum of the log of the ratio of share-weighted relative prices of non-traded goods and the log of the real exchange rate based on the prices of traded goods, $q_t^T = e_t + p_t^{T*} - p_t^T$

$$q_t = (1-g^*) (p_t^{N*} - p_t^{T*}) - (1-g) (p_t^N - p_t^T) + q_t^T, \quad (3)$$

If PPP is taken literally, only q_t^T can be expected to remain constant over long horizons. The long-run real exchange rate based on aggregate price levels q_t , in contrast, may assume any value, depending on the difference between relative non-traded goods prices.

Clearly, when q_t is not constant, not even in the long run, the relationship between the real exchange rate and the real interest differential can no longer be characterized by equation (2). Therefore, in some of the previous studies (Throop 1993, Chinn and Meese 1995, MacDonald 1995, Faruq e 1995) the constant long-run value \bar{q} in (2) was replaced by its counterpart from (3), assuming that the long-run real exchange rate varies with relative non-tradables prices while the real exchange rate based on tradables prices q_t^T remains constant. This results in an augmented version of (2) such as

$$q_t = \bar{q}^T + \frac{1}{1 - \alpha_k} ({}_k r_t - {}_k r_t^*) + (1-g^*) (p_t^{N*} - p_t^{T*}) - (1-g) (p_t^N - p_t^T), \quad (4)$$

which may be simplified to yield

$$q_t^T = \bar{q}^T + \frac{1}{1-\alpha_k} ({}_k r_t - {}_k r_t^*) , \quad (5)$$

since (3) implies $(1-g^*)(p_t^{N*} - p_t^{T*}) - (1-g)(p_t^N - p_t^T) = p_t^* - p_t^{T*} - (p_t - p_t^T)$. In the literature, estimation is typically based on the unrestricted version (5) which is inefficient, given that g and g^* can usually be treated as known.

The more severe problem with (4) and (5), however, is that it remains an *ad hoc* adjustment. It does not take into account the fact that when the aggregate price level-based real exchange rate does not revert to a constant mean in the long run, investors will adjust their expectations accordingly. As a result, the dynamic adjustment pattern of prices and the nominal exchange rate which is the basis of the relationship (2) is no longer represented by equation (1). Below, I will present evidence showing that it is this misspecification which may in part be responsible for the inferior econometric results of earlier studies.

The dynamics assumed in (1), however, now apply to the real exchange rate based on tradables prices, q_t^T , giving $E_t(q_{t+k}^T - \bar{q}^T) = \alpha_k(q_t^T - \bar{q}^T)$. Moreover, with investors aware of PPP only holding on a tradables basis, the real interest parity condition ${}_k r_t - {}_k r_t^* = -E_t(q_{t+k} - q_t)$ has to be corrected for the presence of non-tradables in a similar way. This yields ${}_k r_t^T - {}_k r_t^{T*} = -E_t(q_{t+k}^T - q_t^T)$ where ${}_k r_t^T - {}_k r_t^{T*} = {}_k i_t - {}_k i_t^* - E_t(\Delta_k p_{t+k}^T - \Delta_k p_{t+k}^{T*})$ may be called the real interest rate differential based on tradables prices. Combining the assumption concerning the dynamics with the modified interest parity condition gives

$$q_t^T = \bar{q}^T + \frac{1}{1-\alpha_k} ({}_k r_t^T - {}_k r_t^{T*}), \quad (6)$$

the new reduced form equation of the sticky-price model that applies when PPP holds for traded goods only.

Conveniently from an econometric point of view, equation (6) preserves the simple-regression style form of equation (2) with a constant and the ex ante real interest rate differential as the only regressor to explain the log of the real exchange rate. The assumption that PPP is confined to tradables is taken into account by deflating both the nominal exchange rate and the nominal interest rates by the prices of tradables instead of aggregate price indices. As regards economic interpretation, what distinguishes equation (6) from equation (2) is that since investors expect PPP to hold for traded goods only, they focus exclusively on traded goods prices such as those for manufactures or commodities when calculating whether an investment abroad is more profitable than one at home. Rents, real estate prices or fares for public utilities, in contrast, play no role in their calculation since these prices have no impact on the nominal exchange rate. As

there is no way these prices will interfere with financial arbitrage, they do not enter the UIP-relation and, therefore, do not show up in the reduced form of the sticky-price model.

Engel (1999) has recently shown that the contribution of changes in relative non-traded goods prices to the overall movement of a number of U.S. dollar-based real exchange rates is quite small relative to the contribution of changes in q_t^T ; in most cases, the latter account for more than 90 percent of the real exchange rate's overall variation. From equation (2) and (6) it can easily be shown why this finding does not imply that relative nontraded goods prices are a negligible factor for modeling exchange rate behavior, as one may be tempted to conclude. In fact, the magnitude of the measurable direct influence of changes in relative nontraded goods prices on real exchange rate changes is a poor guide to their overall importance. What counts is the mere existence of changes in relative nontraded goods prices. To remain within Engel's argument: in equation (2) the distinction between traded and nontraded goods not only applies to the real exchange rate but also to the real interest rate differential or, more precisely, to the expected inflation differential, and for the latter the nontraded goods price component is likely to be much more important than it is for the real exchange rate.⁸

2.3 *Current Account Effects*

The current account may interact with the real exchange rate in two ways (Hooper and Morton 1982). First, the current account may be a constraint on the real exchange rate since deviations of the actual current account from its long-run or sustainable level require a change in the "equilibrium real exchange rate", that is the real exchange rate that restores a sustainable current account which, in turn, is the difference between the current and the long-run desired stock of net foreign asset holdings of foreign and domestic investors (Mussa 1984, Faruquée 1995). The actual real exchange rate will move towards the new equilibrium level whenever market participants are convinced that the equilibrium real exchange rate has changed and undertake transactions accordingly. Second, the current account may affect the real exchange rate via its effect on perceived portfolio risk: If assets denominated in a foreign currency provide imperfect insurance against consumption or inflation risks at home, uncovered interest rate parity no longer obtains as risk-averse investors require a premium for holding foreign assets (Adler and

⁸ The reason is that the traded goods component of the expected inflation differential is not affected by the enormous changes in nominal exchange rates which, given the evidence on the failure of the law of one price (see also Engel 1999, appendix C), are likely to dominate the traded goods component of the real exchange rate and reduce the nontraded goods component accordingly.

Dumas 1983).⁹ Sign and magnitude of this risk premium depend on whether the home country is a net creditor or a net debtor and on the share of its net claims or debt, respectively, in world wealth - which in turn depends to a large degree on the historical record of its current account. Changes in the current account will change the country's net creditor/debtor position, change the risk premium and by this affect the real exchange rate. Note that, both the risk premium and the sustainability interpretation imply that the real exchange rate is affected by the stock of net foreign asset holdings, rather than by the flow of international transactions.

Denoting net foreign assets of the domestic country in relation to total wealth as f and assuming a linear relationship between net foreign assets and the real exchange rate, the modified reduced form sticky-price model to be estimated below for the real D-mark/U.S. dollar rate reads

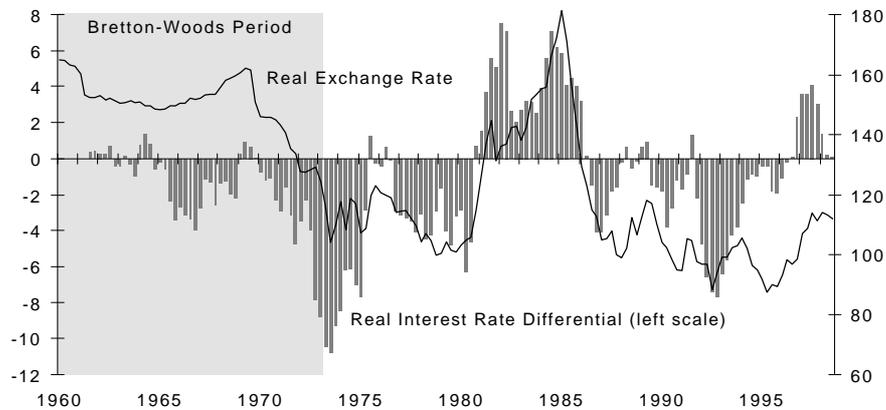
$$q_t^T = b_0 + b_1({}_k r_t^T - {}_k r_t^{T*}) + b_2 f_t + e_t \quad (7)$$

with $b_1 = 1/(1 - q_k) > 1$ since $0 < q_k < 1$. The coefficient $b_0 = \bar{q}^T$ is the PPP-value of the real exchange rate based on tradables prices which, under the risk premium interpretation of the net foreign asset variable, is to be considered as the real exchange rate's constant equilibrium level. Under the interpretation that focuses on the long-run sustainability of the current account, the equilibrium real exchange rate is found by adding the net foreign asset term to the PPP-level, $\bar{q}^T + b_2 f_t$. In this case, the equilibrium value may vary over time with net foreign asset holdings. The coefficient on the stock of net foreign assets b_2 should definitely be negative under the latter interpretation: an increase in net claims to the foreign country makes necessary a real appreciation of the domestic currency that causes net exports to fall and thus restores current account balance. Under the risk premium interpretation, b_2 could be positive as well as negative. Given that the United States have been net debtor to Germany over most of the last two decades, and assuming that holding U.S. dollar denominated assets do indeed provide low consumption insurance to German investors¹⁰, one can conjecture that b_2 would also have to be negative under this interpretation.

⁹ Lewis (1995) and Engel (1996) are recent surveys of the literature on risk premiums in foreign exchange markets.

¹⁰ Income from German and income from U.S. sources may be positively correlated since an appreciation of the dollar against the mark, which would mean a capital gain for a German investor holding U.S. dollar denominated assets, usually comes along with an increase in German net exports that boosts German GDP.

Figure 1:
Quarterly D-mark/U.S. dollar real exchange rate and
U.S.-German real 3-month interest rate differential



Note: Quarterly data as defined in the text, 1960:1-1998:3. Real exchange rate on the basis of producer prices in manufacturing (1990=100). Real interest rate differential in percentage points as defined in the text.

Source: IMF, International Financial Statistics; Deutsche Bundesbank, balance of payments statistics; own calculations.

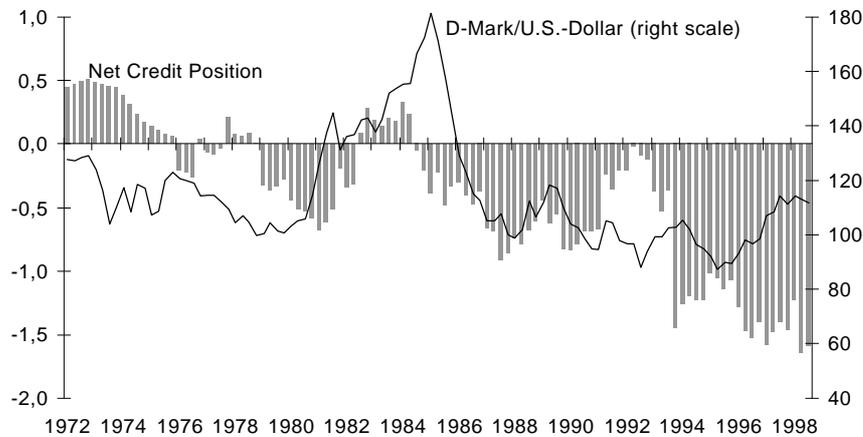
3 Data Description and Preliminary Analysis

Apart from the series on net foreign assets all data come from the International Financial Statistics and is available from the second quarter of 1961 to the third of 1998. As indicated, I use quarterly data, a choice that is primarily dictated by the data used to construct the series on net foreign assets, most of which are not collected at higher frequencies. Availability apart, quarterly data may also be more advantageous than data of higher frequency for sorting out the fundamentals' impact on the real exchange rate as it reduces the variability of the latter without much affecting that of the former.¹¹

The real exchange rate based on the prices of tradable goods, q_t^T , is calculated by taking the logarithm of the average quarterly nominal D-mark/U.S. dollar rate (IFS line 134rf) adding the logarithm of U.S. producer prices in manufacturing and subtracting the logarithm of the corresponding German series (both IFS line 63). To obtain an estimate of the ex ante real interest rate differential I choose $k = 1/4$, calculate the difference between the German 3-month (1 quarter) interbank deposit rate (IFS 134bs) and the U.S. treasury bill rate (IFS 111cs, 111c until 1974:4) and subtract from this the difference between the U.S. and the German expected annualized 3-month rate of change of the index of producer prices in manufacturing (IFS line 63). Following Edison and Pauls

¹¹ Baxter (1994) finds that the correlations between real exchange rates and real interest rates are strongest at trend and business-cycle frequencies while there is no relationship at high frequencies.

Figure 2:
United States net credit position vis-à-vis Germany
and the D-mark/U.S. dollar real exchange rate



Note: Quarterly Data as defined in the text, 1972:1-1998:3. Real exchange rate based on producer prices in manufacturing (1990=100).

Source: IMF, International Financial Statistics; Deutsche Bundesbank, balance of payments statistics; own calculations.

(1993), I approximate price expectations with respect to the following quarter by a centered moving average; some experimentation showed that an average over the rates of change of the last two quarters, the current quarter and the future two quarters¹² provides the best fit.¹³ Figure 1 shows the history of the real exchange rate and the real interest differential defined this way since 1960. Obviously, there is a close positive correlation between the two series which, apparently, even held in the 1960s and early 1970s, that is before the break-up of the Bretton-Woods system of fixed exchange rates.

In contrast to most previous studies, I approximate net foreign asset holdings by genuine stock data. The data is constructed from various series on the financial claims and liabilities of larger German non-bank businesses and banks (plus their U.S. affiliates and subsidiaries)¹⁴ vis-à-vis the United States. The quantitatively most important series has been published in the Deutsche Bundesbank's monthly balance of payments statistics since 1971;¹⁵ they are updated on a quarterly basis. They comprise claims and liabilities

¹² Since no information on prices in 98:4 and 99:1 was available at the time of writing, the last two 'future' observations were generated by use of an exponential smoothing algorithm.

¹³ Note that such a combination of forward-looking and backward-looking price expectations in favor of a purely forward-looking scheme implied by rational expectations is supported by the results of Fuhrer (1997) and Roberts (1998).

¹⁴ Reporting are non-bank companies with balance sheet assets exceeding DM 3 Mio. and banks with claims and liabilities exceeding DM 20 Mio.

¹⁵ The year 1971 refers to the starting point of the series on financial claims and liabilities of banks. Reporting of foreign affiliates and non-banks started in 1980 and 1984, respectively. In the construction of the overall series, the observations of the individual series were simply added when available.

arising from credit relationships of the reporting units; claims and liabilities from securities and foreign direct investments (FDI) are not included. According to German capital accounts statistics, credits made up roughly one third of both capital inflows to and capital outflows from the United States in the mid-1990s, or half of all flows of financial (non-FDI) capital.

In addition, to account for the changes in the flows of securities from Germany to the U. S., I use data from the same Bundesbank statistic to construct a series on net security flows. Moreover, because equation (7) requires the stock data to be put in relation to total wealth, I divide the series on the net stock of credit by U.S. nominal seasonal adjusted GDP (converted into D-marks by PPP-exchange rates) and the series on the net flow of securities towards the United States by the corresponding quarter-to-quarter change of the same GDP series. The resulting series are denoted f_t^c (net credit stock) and Δf_t^s (net flow of securities) in the following. The history of $(-f_t^c)$ is shown in Figure 2 along with the D-mark/U.S. dollar real exchange rate based on tradables prices. In line with the increasing current account deficits of the United States against Germany, the German net credit position vis-à-vis the U.S. has been rising for most of the time since the late 1970s.

To determine whether the time series are stationary, I employ standard Augmented Dickey Fuller (ADF) tests for non-stationarity, a modified higher power version of the ADF test proposed by Elliot et al. (1996) and the KPSS test for stationarity (Kwiatkowski et al. 1992). The sample period covers the modern floating period (1973:2-1998:3). According to the results given in Table 1, there are two stationary and two non-stationary variables in the system. The ex ante real interest differential, ${}_k r_t^T - {}_k r_t^{T*}$, and the net capital outflows in the form of securities, Δf_t^s , can safely be treated as stationary: the ADF tests reject the null hypothesis of non-stationarity whereas the KPSS test does not reject the null of stationarity. The real exchange rate q_t^T and the net foreign asset position, f_t^c , in contrast, are both I(1) in this sample; non-stationarity and non-trend stationarity are not rejected whereas stationarity, trend-stationarity and non-difference stationarity are rejected.

Having determined the time series properties of the individual variables, the next step in the preliminary data analysis would usually be to find out how many cointegration vectors there are in the present four-variable system by use of Johansen's (1995) maximum-likelihood framework. However, given that only two variables are non-stationary,

Table 1:
Tests for non-stationarity and stationarity

	q_t^T	${}_k r_t^T - {}_k r_t^{T*}$	f_t^c	Δf_t^s
	<i>Levels</i>			
<i>ADF(c, t)</i>	2.49	-	-2.54	-
<i>ADF(c)</i>	-2.24	-	-1.50	-
<i>ADF</i>	-0.03	-2.61**	-0.45	-6.54**
<i>ADF-GLS(c, t)</i>	-0.17	-	-0.32	-
<i>ADF-GLS(c)</i>	-0.20	-2.65**	0.03	-6.60**
<i>KPSS(c, t)</i>	0.23**	-	0.13*	-
<i>KPSS(c)</i>	0.53**	0.33	1.52**	0.11
	<i>First Differences</i>			
<i>DF</i>	-4.23**	-4.38**	-11.92	-10.82

Notes: Sample period 1973:2-1998:3 (T = 102). *ADF(c,t)* indicates an Augmented Dickey Fuller Test for non-stationarity with constant (c) and deterministic trend (c, t). *ADF(c)* the counterpart without a deterministic trend, and so on. *ADF-GLS* stands for the modified *ADF* test proposed by Elliot et al. (1996) based on local demeaning ($\bar{c} = -7$) and detrending ($\bar{c} = -13.5$). In each regression, the lag length was chosen such that the residuals were serially uncorrelated. KPSS is the test for stationarity proposed by Kwiatkowski et al. (1992) where a lag truncation parameter of 4 for the non-parametric correction is used to obtain a consistent estimate of the variance. In each case * and ** indicate significance at the 10% and 5% level, respectively. Critical values for the *ADF* and *ADF-GLS(c)* tests are taken from MacKinnon (1991), for the *ADF-GLS(c, t)* from Elliot et al. (1995) and for the KPSS test from Kwiatkowski et al. (1992).

there can at maximum be one cointegration relationship among these variables. The hypothesis that this relationship does indeed exist can be tested more conveniently within a conditional single-equation model, the construction of which I will turn to next.

4 The Empirical Model

4.1 Econometric Approach

Let x_t be a vector comprising the real exchange rate and the variables suspected to influence it, that is $x_t = (\bar{q}_t^T, r_t - r_t^*, f_t^c)$, which follows a vector autoregressive process of order p that may be expressed as the vector error correction model (VECM)

$$\Delta x_t = A x_{t-1} + \sum_{j=1}^{p-1} A_j \Delta x_{t-j} + D_t + e_t, \quad t = 1, \dots, T. \quad (8)$$

where e_t is an independent $N(0, \Sigma)$ sequence, A and A_j are parameter matrices and D_t a vector of deterministic components such as an intercept or a 0/1-dummy variable. In case that some of the variables in x_t are cointegrated matrix A has reduced rank and can therefore be expressed as $A = ab$, where a contains the adjustment parameters and b is the vector of long-run parameters on which interest will center in the following.

To obtain a conditional model from (8), partition x_t into $x_t = (y_t', z_t')$, where y_t contains the endogenous variable q_t^T and the vector z_t consists of the variables q_t^T is conditioned upon, i.e. $r_t^T - r_t^{T*}$ and f_t^c . Equation (7) can then be rewritten as

$$\Delta y_t = a_{1,1} b_1' y_{t-1} + a_{1,2} b_2' z_{t-1} + A_0 \Delta z_t + \sum_{j=1}^{p-1} A_{1j} \Delta x_{t-j} + D_{1t} + e_{1t}, \quad (8a)$$

$$\Delta z_t = a_{2,1} b_1' y_{t-1} + a_{2,2} b_2' z_{t-1} + \sum_{j=1}^{p-1} A_{2j} \Delta x_{t-j} + D_{2t} + e_{2t} \quad (8b)$$

where $t = 1, \dots, T$ and $(e_{1t}, e_{2t})' \sim N(0, \Sigma = \begin{pmatrix} \Sigma_{e_1 e_1} & \Sigma_{e_1 e_2} \\ \Sigma_{e_2 e_1} & \Sigma_{e_2 e_2} \end{pmatrix})$.

Since it is known that there can at most be a single cointegration vector, the analysis of this system can proceed in the three steps proposed by Boswijk (1994, 1995). Assuming weak exogeneity holds for the conditioning variables, the cointegration vector b can be estimated consistently and efficiently from equation (8a) by OLS as proposed by Stock (1987). Nothing can be learned from the other equations in the system. The existence of cointegration can be analyzed by testing for the joint exclusion of the variables in levels ($H_0: a_{1,1} b_1' = 0, a_{1,2} b_2' = 0$) by means of a Wald test. Critical values for the test statistic, denoted χ_c below, are given in Boswijk (1994). If the null hypothesis of no cointegration is rejected, the assumption of weak exogeneity can be tested by adding the

estimated equilibrium error $\hat{v}_t = y_t - \hat{b}z_t$ to the marginal model (8b) and testing its joint or individual significance in the various equations of (8b) by means of LM and Wald tests for $H_0: a_2 = 0$ (Johansen 1992, Boswijk 1995). These test statistics, which have standard distributions given cointegration, will be denoted c_{exo}^2 and t_{r-r^*} and t_{f^c} , respectively, below. In case weak exogeneity is rejected, the OLS estimator of the cointegration vector remains consistent but loses its efficiency property.¹⁶

Transformed into a conditional ECM, where now $y_t = q_t^T$ and $z_t = (r_t^T - r_t^{T*}, f_t^c)$, equation (7) reads

$$\begin{aligned} \Delta q_t^T = & a_0 + a_1 q_{t-1}^T + a_2 (r^T - r^{T*})_{t-1} + a_3 f_{t-1}^c + \sum_{j=1}^{p_1-1} a_{1j} \Delta q_{t-j}^T \\ & + \sum_{j=0}^{p_2-1} a_{2j} \Delta (r^T - r^{T*})_{t-j} + \sum_{j=0}^{p_4-1} a_{3j} \Delta f_{t-j}^c + \sum_{j=0}^{p_5-1} a_{4j} \Delta f_{t-j}^s + u_t \end{aligned} \quad (8)$$

In this model $a_0/a_1 = b_0 = \bar{q}_t^T$, $a_2/a_1 = b_2$ and $a_3/a_1 = b_3$, implying that $b_1 > 1$ should be found to hold if equation (7) is a correct description of the long-run relationship between the variables. For inference on the long-run coefficients b_0, b_1, b_2 , their standard errors can be estimated directly from the Bewley-transformed regression (Wickens and Breusch 1988). Note that the series representing the net *flow* of security investments Δf_t^s does not enter the long-run relationship in this specification, which according to (6) is confined to the relation between the *stock* of net foreign assets and the real exchange rate. The flow variable does, however, show up in the short-run dynamics, together with the first differences of the long-run variables.

4.2 Results of the Estimation

Estimation of (8) is based on a sample running from 1973:2- p to 1998:3. This sample choice requires the use of p observations from before the break-up of the Bretton-Woods system. At first sight, this may look unsatisfactory as mixing data from different policy regimes may be suspected to induce instability of the parameter estimates. A natural alternative would be to take 1973:2 + p as the beginning of the sample. However, some experimentation with regressions that exclusively used post-Bretton-Woods observations not only showed higher standard errors but also more unstable coefficient

¹⁶ A number of Monte Carlo studies (i.e. Phillips and Loretan 1991, Inder 1993, Stock and Watson 1993, Boswijk 1995) indicate that when weak exogeneity fails to hold, estimates of long-run parameters from partial models can be biased in small samples. The fact that in the present study $r^T - r^{T*}$ was found to be stationary may aggravate the bias problem in this case.

estimates and significantly worse performance in out-of-sample forecasting, so including some observations from the Bretton-Woods period seems preferable.¹⁷

In the exploratory phase of the modeling process I employ residual based misspecification tests — such as tests for first- and higher order autocorrelation, non-normality, conventional and autoregressive heteroscedasticity (ARCH) of first and fourth order and tests for general misspecification (see e. g. Harvey 1990, ch. 5) — to judge the adequacy of the model's specification. These tests and the Schwarz-criterion indicate that a lag order of $p = 5$ gives a sufficient representation of the dynamics. The test for non-normality and a search procedure based on studentized residuals point to the existence of two poorly modeled observations ("outliers") in 1974:1 and 1974:3 which are accounted for by 0/1-impulse dummies (denoted I). To achieve a more parsimonious and thus more stable representation of the data, the model is then simplified by deleting subsequently all of the first differenced variables with the lowest t -ratios and imposing equality constraints on some of the remaining coefficients.

The result is the following equation for the change of the real exchange rate with only 10 parameters; the estimated long-run coefficients b_i — taken from the Bewley-transformed regression — are presented in the error correction term in squared brackets and t -ratios are given in parentheses beneath the coefficient estimates:

$$\begin{aligned} \Delta q_t^T = & -0.318 \left[4.831 - 3.922(r^T - r^{T*})_{t-1} + 0.191f_{t-1}^c \right] & (10) \\ & (8.0) \quad (326.5) \quad (11.9) & (8.8) \\ & + 0.150(\Delta q_{t-1}^T + \Delta q_{t-3}^T + \Delta q_{t-4}^T) + (0.318 \cdot 3.922)\Delta(r^T - r^{T*})_t + 0.872\Delta(r^T - r^{T*})_{t-4} \\ & (2.8) & (7.7) & (3.9) \\ & + (0.318 \cdot 0.191)(\Delta f_{t-1}^c + \Delta f_{t-2}^c) + 0.006\Delta f_{t-3}^S + 0.013\Delta f_{t-4}^S \\ & (5.7) & (1.3) & (3.6) \\ & + 0.096(I74:1 + I74:3) \\ & (3.7) \end{aligned}$$

R^2 -ECM = 0.58; R^2 -Bewley = 0.59; $s = 0.034$, $T = 102$ [1973:2-1998:3],
 DW: 2.05, LM(1): 0.869, LM(4): 0.774, LM(8): 0.816,
 JB: 0.893, BP: 0.591, ARCH(4): 0.438, RESET: 0.971
 ADF: -3.817 [3.74], x_m : 61.52 [15.22]
 $c_{exo}^2(2)$: 4.711 [5.99], $t_{r^T - r^{T*}}$: -1.824, t_{f^c} : -1.412

¹⁷ The reason for the superior estimates obtained from the longer sample may be that without the Bretton-Woods observations and a lag order of $p = 5$, estimation starts with 1974:3 while by starting estimation with 1973:2, the sampling period is effectively extended back to 1972:1 which may have the advantage of including a larger part of the cycle the real interest rate differential was in at that time, as can be seen from Figure 1.

From a statistical point of view, the estimated model seems to be well specified, with tests showing no signs of residual autocorrelation, heteroscedasticity or non-normality.¹⁸ The results of both the *ADF* test for cointegration and Boswijk's (1994) χ^2 -tests indicate that the variables are likely to be cointegrated. The *ADF* test rejects the hypothesis that the equilibrium error $\bar{v}_t = q_t^T - \bar{\beta}_2(r_t^T - r_t^{T*}) - \bar{\beta}_3 f_t^c$ is non-stationary at the 5% level with a t -value of -3.82. The Wald test for the joint exclusion of q_{t-1}^T and f_{t-1}^c from the regression produces a value of 61.52 which is far above the relevant 1%-critical value of 15.22 given in Boswijk (1994, Table B.2).¹⁹ Moreover, the high t -ratio of the coefficient of the equilibrium correction term also indicates cointegration.

The results of the tests for weak exogeneity are somewhat less comforting. The null hypothesis that the equilibrium error enters both marginal equations (Boswijk 1995) is rejected at the 5%- but not on the 10%-level of significance ($C^2(2) = 7.41$). Judged from the t -ratios of the constructed equilibrium errors $\bar{\Pi}_t$ in the marginal equations (8b) weak exogeneity is obviously a reasonable assumption for the current account variable f^c . For the real interest rate differential weak exogeneity is rejected at the 10 but not at the 5 percent level of significance.

¹⁸ DW is the Durbin-Watson statistic, $LM(k)$, $k = 1, 4, 8$ are Lagrange multiplier tests for autocorrelation based on 1, 4 and 8 lags, respectively, JB the Jarque-Bera test for residual non-normality, BP the Breusch-Pagan test for heteroscedasticity, ARCH(4) an LM test for fourth order autoregressive conditional heteroscedasticity, RESET the test for a general misspecification of the model (see e. g. Harvey (1990) for details on these tests). DW apart, the numbers indicate the marginal significance levels of the calculated test statistics. None of tests indicates a misspecification of the equation.

¹⁹ The test implies that the constant term is not restricted to the cointegration space which seems to be the more plausible hypothesis given the trending behavior of f^c .

Table 2:
Long-run relation: estimates from different methods

	intercept	$r^T - r^{T*}$	f^C
Conditional ECM	4.831 (326.5)	+3.922 (11.9)	-0.191 (8.8)
Johansen (1991)	4.821 (370.8)	+3.415 (11.7)	-0.199 (17.5)
Stock/Watson (1993)	4.858 (612.8)	+4.274 (20.2)	-0.237 (18.9)

Notes: Sample period 1973:2-1998:3. Numbers in parenthesis are t -ratios. For the Johansen (1991) procedure the lag order was 6. A total of 10 impulse dummies (73:1, 73:2, 75:1, 80:2, 80:4, 82:3, 91:1, 91:2, 93:4, 97:2, 97:3, 98:2) were needed in the VAR to obtain a normally distributed residual vector. The net foreign asset variable was assumed to be weakly exogenous for the remaining variables in the system.

For the DGLS procedure of Stock and Watson (1993) procedure 4 leads and lags of the first differenced conditional variables were used. The computation of the adjusted covariance matrix was based on an AR(5) process for the residuals.

Still, since inference on the cointegration parameters may be distorted if weak exogeneity does not hold for the conditioning variables, I reestimate the complete system (8) by the Johansen (1991) procedure with a lag order of 6 and the same set of impulse dummies as before to achieve an approximately normal residual distribution.²⁰ In addition and as a check-test, I calculate the DGLS estimator proposed by Stock and Watson (1993) which is also efficient asymptotically and does not require the estimation of the complete system and, therefore, makes do without the complete set of impulse dummies. Here, I choose 4 leads and 4 lags of the first differences of the right-hand side variables to account for the endogeneity of the real interest rate differential. The results of the two procedures are shown in Table 2 together with the estimates from the conditional ECM. With both methods, the estimates turn out to be significantly different from zero, confirming the results of the single-equation analysis. Moreover, the parameter estimates from all three methods are quite similar, confirming the robustness of the conditional analysis and justifying the decision to work with the conditional model in the following.

²⁰ The restriction that f^s does not enter the cointegration relationship is not rejected. In contrast to the tests conducted on the basis of the estimates from the conditional model, however, I find both the net foreign asset variable and the real interest rate differential to be weakly exogenous for the real exchange rate. The reason may be that the test I used for the conditional model is more efficient since I allowed the dynamics and the impulse dummies to be equation specific by eliminating all insignificant lagged variables and impulse dummies and estimating the model as a system of seemingly unrelated regressions.

4.3 Long-Run and Short-Run Properties of the Model

The long-run properties of the model are given by its static long-run solution given in squared brackets in equation (10) and in Table 2. Both long-run coefficients have theoretically sensible signs and magnitudes: According to the estimates, the D-mark appreciates in real terms vis-à-vis the dollar both as a result of an increase of the German over the U.S. real interest rate and in reaction to an increase in German claims vis-à-vis the United States. The fact that changes in the net foreign asset stock effectively influence the real exchange rate with a lag of three quarters supports the view that this effect represents the markets expectations concerning the long-run sustainability of a current account imbalance rather than a risk premium. In the perception of the market, the long-run, equilibrium real exchange rate is not a constant but varies with stock of foreign claims or liabilities, following $q^T = \bar{q} - 0.191f^C$.

The estimated coefficient of the real interest rate differential is significantly above unity which is in line with the stylized Dornbusch-Frankel model (5).²¹ All else equal, a difference between the German and the American annualized 3-month real interest rate of 1 percentage point in favor of Germany will have appreciated the D-mark vis-à-vis the dollar by roughly 4 percent in real terms after all adjustments have taken place. By implication of the dynamics assumed for the real exchange rate when deriving equation (6), q_t^T follows a first-order autoregressive process with an annualized speed-of-adjustment parameter $\alpha_{1/4}$ of 0.75, thus a deviation from exchange rate equilibrium caused by a non-zero real interest rate differential has an average half-life of 2.4 years. This estimate is somewhat smaller than the 3 to 4 years usually found in studies that focus on (aggregate price level-based) PPP as the equilibrium concept (see Froot and Rogoff 1995 and Rogoff 1996). The divergence may be explained by the different concepts for exchange rate equilibrium: PPP-studies based on tests for mean-reversion of the real exchange rate by construction treat all changes in the real exchange rate as deviations from equilibrium, including those caused by changes in non-tradables prices or net foreign asset stocks, which in the present study would be counted as changes in the equilibrium rate. Given the fact that there have been quite some changes in the latter two variables over the sampling period, PPP-studies should measure stronger and longer lasting deviations from equilibrium than are measured here as long as the effects of those changes on the real exchange rate have not compensated each other exactly.

To facilitate the interpretation of the model's dynamic properties, it is helpful to rewrite equation (10) as an autoregressive distributed lag model in levels. The result is

²¹ The hypothesis that $b_2 = 1$ is rejected with a t -ratio of 9.856.

$$\begin{aligned}
q_t^T - \bar{q}^T = & 0.812q_{t-1}^T - 0.142(q_{t-2}^T - q_{t-3}^T + q_{t-4}^T) + 1.251(r^T - r^{T*})_t \\
& - 0.958[(r^T - r^{T*})_{t-4} - (r^T - r^{T*})_{t-5}] - 0.063f_{t-3}^C - 0.006(f_{t-3}^S + f_{t-4}^S) + 0.013f_{t-5}^S
\end{aligned} \tag{11}$$

where dummies are not shown for simplicity. An equivalent representation of the dynamics is given by means of the cumulated values of the normalized lag weights $\mathbb{1}y_t/\mathbb{1}x_{t-i}$, $i = 0, 1, 2, \dots$, $y = q^T$, $x = (r^T - r^{T*}, f^c)$ shown in Figure 3.

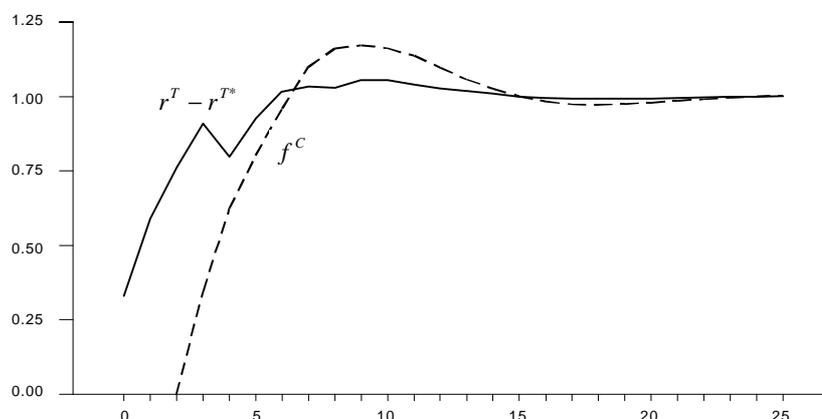
Both equation (11) and Figure 3 show that real interest rates affect the real exchange rate contemporaneously as is predicted by the Dornbusch-Frankel model. Moreover, the magnitude of the impact-interest rate semi-elasticity estimate of 1.25 is also in accordance with that model: Assuming that goods prices and inflation expectations remain unchanged in the quarter of the shock, such that only nominal variables change, the estimate implies that an expansionary monetary policy, say in Germany, which increases the nominal (and real) interest rate differential in favor of the United States by one percentage point, instantaneously appreciates the dollar by somewhat more than one percent. This is exactly the appreciation needed to restore uncovered interest parity as it implies that the nominal exchange rate overshoots its new long-run value²² and thus generates the expectation of a future depreciation of the dollar by one percent. Therefore, if the assumption of price stickiness in the first quarter is correct, the estimate can be interpreted as evidence in favor of the overshooting hypothesis.

Another dynamic property of interest is the presence of the autoregressive coefficients q_{t-i}^T , $i = 1, \dots, 4$ which generate long solved distributed lags of the real interest rate differential and the foreign assets variable. The real exchange rate, thus, adjusts slowly to shocks as can be seen from the graphs of the cumulative normalized lag weights showing that a once-and-for-all increase in both the real interest differential or the net foreign asset stock take about 10 quarters to exert their full effect on the real exchange rate. This implies that shocks to the real interest rate differential need only be moderately persistent (that is, the autocorrelation coefficient of the real interest rate differential has to be larger than $1 - 0.812 = 0.188$) to generate real exchange rate dynamics of the type reported by Eichenbaum and Evans (1995). Using impulse response analysis, these authors find the adjustment of real exchange rates to a monetary shock to proceed much more gradually than the Dornbusch-Frankel framework would imply²³ and discuss noise-trading behavior as a possible explanation.

²² The point estimate of 1.25 for the impact interest elasticity would - under the assumptions made above- imply that the market expects the nominal exchange rate to appreciate by 0.25 percent in the long-run after the interest rate shock.

²³ See Clarida and Gali (1994) for similar results.

Figure 3:
Dynamic properties of the model:
cumulative normalized lag weights for $r^T - r^{T*}$ and f^C



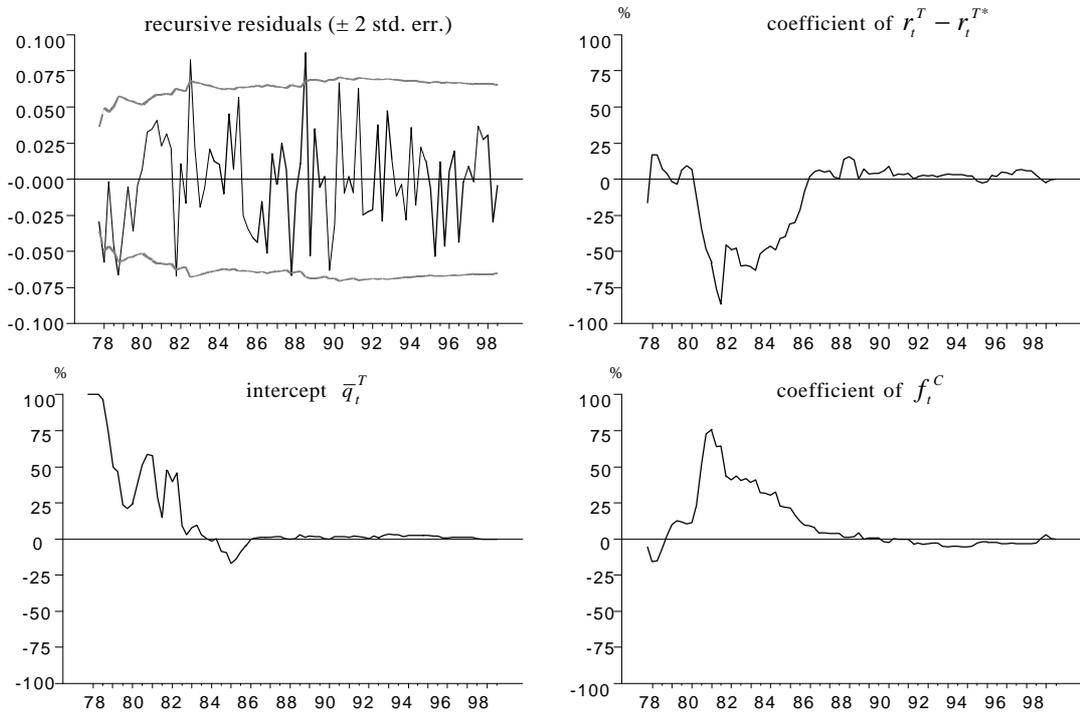
Overall, equation (10) explains 58 percent of the in-sample variation of the change of the real exchange rate; the corresponding figure for the level of the real exchange rate - again calculated from the Bewley-transformed regression - is 59 percent. In view of the substantial volatility of the real exchange rate both figures can be considered high. The graphical comparison of the actual to fitted values of the model (given below in Figure 5 for the sample that includes Bretton-Woods observations) supports the assessment of a good in-sample fit. Note that the fundamentals-based model is capable of explaining most of the dollar's strong real appreciation vis-à-vis the mark during the first half of the 1980s. This movement is often taken as an example of a typical financial market "bubble", an exchange rate change that is largely unrelated to the underlying fundamentals and which, due to its severe effects on the real economy, may be a reason to "throw sand in the wheels of financial markets". According to the model and the evidence on real interest rates at that time, as presented in Figure 1, in contrast, most of the appreciation can be attributed to the exceptionally high real interest rate differential in favor of the United States which was caused by opposite monetary and fiscal policies on both sides of the Atlantic²⁴.

4.4 Stability Analysis

Before moving to the analysis of the model's forecast performance, which is basically a check of its stability, note that other formal tests for a structural break or for parameter changes do not indicate any problems with this point; good forecast results can therefore be expected. As a first informal check of the stability of the model, consider the

²⁴ See Branson (1988) and Obstfeld (1988) for a discussion.

Figure 4:
One-step residuals and percentage deviations of recursively estimated long-run coefficients from their full-sample estimate



results from a recursive reestimation of equation (10) depicted in Figure 4. The statistics are one-step residuals plus or minus twice the full-sample standard error and the percentage deviations of the estimated long-run coefficients from the full-sample estimates as given in equation (10). Overall, the figure supports the view of a stable relation in the period of interest, although the one-step residuals indicate some instability in the early and the late 1980s which probably reflect the model initial difficulties to cope with the strong dollar appreciation of the mid 1980s. This also shows up in the recursive estimates of the long-run coefficients which show quite some variation in the early 1980s but remain remarkably close to their full-sample estimate in the second half of the sample.

More formally, a standard Chow-test rejects the hypothesis that there has been a structural break during the last 20 observations of the sample, with $F(20,90) = 0.879$, receptively. The assessment is the same when the tests are repeated for the last ten observations of the sample. It is well known, however, that tests that require the a priori selection of a breakpoint have low power when the break occurs at some other point of the sample. Tests which simultaneously determine the most likely date of the break and test its significance are more appropriate in this case. Unfortunately, the distributional

theory for most of the tests is only developed for models with stationary variables.²⁵ Therefore, I reestimate equation (10) with the partly non-stationary variables in levels replaced by the lagged estimated equilibrium error $\bar{\mu}_{t-1}$ which is stationary by virtue of cointegration. The results of the tests applied to the modified version of equation (10) are presented in Table 4. The tests are the CUSUM test of Brown, Durbin and Evans (1975) and its version based on OLS-residuals by Ploberger and Krämer (1992), the Quandt (1960) likelihood ratio (QLR) with critical values from Andrews (1993) and the Nyblom (1989) test for random walk parameters in the version of Hansen (1992a). None of the tests indicates an instability of the model. Most tellingly, the sequence of break-point- F -tests that makes up the QLR reaches its maximum in the first quarter of 1980 but the value of the test statistic of 7.98 for that period is far from the critical value of 25.47 tabulated by Andrews (1993, Table I, for a model with 9 parameters and a trimming region of $\pm 15\%$ of the sample). Similarly, the L -statistic is insignificant, rejecting the hypothesis that the coefficients follow a random walk.²⁶

Table 3:
Tests for structural break and parameter instability

Test	Statistic	5%-crit. value
CUSUM:	0.357	[0.948]
OLS-CUSUM:	0.511	[1.360]
QLR:	10.54	[25.47]
L_c	1.80	[2.54]

Notes: Cusum is the test based on recursive residuals by Brown, Durbin and Evans (1975); critical values taken from Harvey (1990, p. 366). OLS-Cusum is the version of the Cusum test based on the standard OLS-residuals as proposed by Ploberger and Krämer (1992), QLR the Quandt likelihood ratio with critical values from Andrews (1993) and a symmetric trimming region of ± 15 percent. L_c is the test for parameter instability of Hansen (1992a) applied to the parameter vector of the complete model, including the variance.

²⁵ See Hansen (1992) for an exception.

²⁶ Supplementary evidence on the stability of the long-run relationship, moreover, comes from the significance of the ADF test for cointegration. Monte Carlo simulations by Gregory et al. (1996) show that the rejection frequency of the ADF test for cointegration decreases substantially if there is a structural break in the cointegration relationship.

5 Encompassing the Random Walk and other Rival Specifications

5.1 Outperforming the Random Walk

Since Meese and Rogoff (1983) showed that the out-of-sample forecasts produced by econometric exchange rate models were easily out-performed by a naive random walk model, beating the random walk forecast has become the standard metric by which to judge empirical exchange rate models. To replicate their exercise, I initially estimate equation (10) through 1992:2 and generate forecasts for the level of the real exchange rate from there on for the next 1, 2, 3, 4, 8 and 12 quarters. I then take one more observation (1993:1) into the sample, reestimate the model and generate a new set of forecasts. Proceeding this way until 1998:3, I get a total of 25 one-step-ahead, 24 two-step-ahead, 23 three-step-ahead and, among others, 13 twelve-step-ahead forecasts for which I compute root means squared forecast errors (RMSE).²⁷

The first row of Table 4 presents the results of the forecast performance exercise for equation (10). For each of the six forecast horizons running from one quarter to three years the table gives Theil's inequality coefficient (U-ratio), that is the forecast RMSE of the estimated model to the forecast RMSE of the random walk. Evidently, the ECM by far outperforms the no-change forecasts of the random walk model over all horizons. Relative forecast performance increases with the forecast horizon and reaches its maximum at two years (8 quarters), where the model's RMSE is merely a quarter of that of the random walk. This is far lower than the results previously reported for structural exchange rate models (Chinn and Meese 1995, Mark and Choi 1997, MacDonald and Marsh 1997) although the results are not directly comparable as these studies used different forecast samples.²⁸ Moreover, even when it comes to predicting the real exchange rate in the following quarter, a discipline many structural empirical exchange rate models fail to outperform the random walk in, equation (10) produces better forecasts: the RMSE of the structural model is less than two thirds that of the random walk. Overall,

²⁷ A visual impression of the in- and out-of-sample predictions of the model estimated on the sample that includes Bretton-Woods observations is given in Figure 5 below. The figure for the model estimated on the floating rates data would look virtually identical.

²⁸ It might be interesting to compare the performance of the present model directly to results of the most recent studies by Mark and Choi (1997) using exactly the same estimation and forecast periods as those authors. Mark and Choi (1997) estimate various real exchange rate models on the basis of monthly data from 1973:3 to 1980:12 and generate 48-step-ahead forecasts for the period of 1985:01 to 1993:11. For the quarterly data used in the present study this translates into an estimation period using data from 1973:2 through 80:4 and a forecast period extending from 1985:1 to 1993:4 for which 16-step-ahead forecasts are generated. The lowest U-ratio Mark and Choi (1997, Table 2) report for their D-Mark/Dollar models is 1.158 whereas for the present model the U-ratio is 0.34. Starting the estimation period in 1961:1, Mark and Choi can improve the minimum inequality coefficient to 0.783. For the present model, extending the estimation period to 1963:1 (see next subsection) also lowered the U-ratio considerably to 0.26.

the results strongly support the view that the macroeconomic fundamentals focused on here can successfully be used to forecast the real D-mark/U.S. dollar exchange rate.

5.2 Rival Specifications and Different Sampling Periods

To assess the impact of different data and model specifications on the in- and out-of-sample performance of the model, I estimate a number of rival specifications and calculate the relevant statistics. As measures of the in-sample fit, I report the R^2 of the ECM, the ADF test for cointegration and the QLR statistics. Out-of-sample forecast performance for the period 1992:4 to 1998:3 is again assessed relative to the random walk model; in comparing the various specifications, the forecast-RMSE of the naive model works as a 'numéraire'. Each rival differs from the base equation (10) in one or more respects as for as modeling of the long-run relationship (7) is concerned. The differences may be a variation of the data (e. g. consumer prices instead of tradables prices), a coefficient restriction (e. g. $b_2 = 0$) or both or a variation of the sampling period. To give the rivals a fair chance, each long-run specification is subjected to exactly the same dynamic modeling procedure that produced equation (10). Both the short-run dynamics of the rival models as represented by the number of lagged variables in first differences and the dummy variables which again are used to account for poorly modeled observations may turn out to be different from (10). In practice, the lag distributions turned out to be identical to the base equation in each case but impulse dummies were employed at different dates.

The results of the exercise are summarized in Table 4. The rivals are divided according to which of the three variables in the long-run relationship has been changed in comparison to equation (10). The specifications under *I.* refer to models with modified definitions of the real interest rate differential. *I.a* varies the way price expectations are modeled. Expected changes in producer prices in manufacturing are approximated by simple current quarter changes instead of the five-quarter moving average of past and future changes used in (10). Although the in-sample fit, as measured by R^2 , falls significantly, the U-ratios increase only by 0.1 vis-à-vis those of equation (10). I conclude that the precise way of approximating price expectations is of minor importance in the present context. However, when consumer prices are used to deflate nominal interest rates instead of producer prices in manufacturing (*I.b*) (and expectations are again modeled by five-quarter moving averages), relative forecast performance worsens considerably. This result is of special interest since this specification is based on equation (6), the unrestricted version of which some of the earlier studies used to account for the effects of relative non-tradables prices on the real exchange rate. According to the results here, this model can be rejected in favor of model (10).

Table 4:
Alternative specifications of the ECM: in-sample fit and
out-of-sample forecast performance in relation to the random walk

Specification	U-Ratio Over ... Forecast Steps						In-Sample Fit		
	1	2	3	4	8	12	R^2	ADF	QLR
eq. (10):	0.66	0.53	0.43	0.34	0.25	0.29	0.59	-5.99*	10.54
<i>I. alternative real interest rate differentials</i>									
I.a: (price expectations)	0.76	0.55	0.45	0.42	0.32	0.36	0.28	-6.52*	13.76
I.b: (consumer prices)	0.89	0.81	0.74	0.69	0.67	0.85	0.29	-4.28*	25.92*
<i>II. alternative net foreign asset variables</i>									
II.a: (no f -variable at all)	0.94	0.96	0.94	0.88	0.89	1.07	0.40	-4.50*	13.09
II.b: (cum.curr.accounts)	0.82	0.82	0.74	0.62	0.37	0.35	0.55	-5.16*	9.06
<i>III. CPI-based models</i>									
III.a: (no f -variable at all)	1.05	1.02	1.00	0.96	0.96	1.15	0.22	-2.38	19.41*
III.b: (cum. curr. accounts)	0.98	0.96	0.90	0.82	0.73	0.86	0.30	-2.61	16.93*
III.c: (as in eq. (10))	0.95	0.97	0.90	0.84	0.80	0.99	0.35	-4.33*	22.21*
<i>IV. variation of sampling/forecast periods</i>									
eq. (10): 1973:2-1986:1, forecasts: 1980:1-1986:1	1.98	1.60	1.33	1.15	0.78	0.68	0.68	-3.65*	38.09*

Notes: Forecast results are based on rolling regressions for the period 1992:3-1998:3 except for results in the last row. U is the ratio of the root mean square forecast error (RMSE) for the level of the log real exchange rate of the estimated model to the RMSE of the random walk model. In-sample results refer to the full-sample (1973:2-1998:3) except for the last row which refer to the sampling period 1973:2-1986:1.

The specifications under *II.* use alternative variables to approximate the net foreign asset position. Line *II.a* reports the performance of a pure real interest differentials based-model where net foreign asset effects are omitted completely. While the in-sample fit of this model is still acceptable, its forecasting performance, particularly over longer horizons, is quite poor in comparison to the base model. As noted above, a number of studies have used cumulated bilateral current account deficits to approximate net foreign asset positions. Line *II.b* shows the results of this modification.²⁹ The in-sample fit and out-of-sample forecast performance is considerably better than that of *II.a*. However, most of this improvement was due to a step dummy which was significant in 85:2 (t -value 4.47). Without that dummy, the results (not reported) strongly resembled those of *II.a*.

²⁹ Like the net foreign asset variables f^e and f^f used in (11), the cumulated current account deficit was divided by seasonal adjusted nominal U.S. GDP.

The three rows above the last row of Table 3 report the results for the specifications usually used in the literature. Both real exchange rates and real interest differentials are based on aggregate price indices such as CPIs (see e. g. Campbell and Clarida 1987, Meese and Rogoff 1988, Edison and Pauls 1993, Baxter 1994). The first is again the pure real exchange rate-real interest rate model without a net foreign asset variable. Both in- and out-of-sample fit of this model are very poor; there is no significant forecast improvement over the naive model, the variables lack cointegration and the parameter estimates show signs of instability. The latter two problems also plague the version of the model that uses the cumulated bilateral current account to approximate the net foreign asset position (*III.b*). Forecasts of this model, nevertheless, are considerably better than those of the previous specification but still far weaker than those of the base model (10). Astonishingly, use of the net foreign asset variable proposed in the present paper, f_t^C , worsens rather than improves forecast performance of the CPI-based model (*III.c*) although non-cointegration is now rejected and R^2 is comparatively high. Overall, the performance of the traditional, CPI-based specifications is very weak in comparison to equation (10).

The last row shows the impact of the sampling period on the results by giving the statistics that turn out when equation (10) is estimated on the sample that ends in 1986:1. This is the sample that was available to Meese and Rogoff (1988) and they estimated their models through the end of 1979 and generated rolling regression forecasts from early 1980 to early 1986. Repeating their analysis on the basis of equation (10) I find that its forecasting performance deteriorates sharply relative to the results for the later forecast period. At forecast horizons up to four quarters the random walk significantly outperforms the model as was found by Meese and Rogoff for their models. Unlike those models, however, equation (10) manages to beat the random walk forecast at long horizons in this sample.

Summing up, the analysis shows that the empirical exchange rate model presented in this paper is capable of encompassing a wide range of alternative specifications, some of which have been used extensively in the literature. Its superior performance hinges on all parts of its specification apart from the particular scheme used to approximate inflation expectations. Still, of the two innovations to the modeling of exchange rate behavior put forth here — genuine stock data to approximate net foreign assets and the consistent treatment of nontraded goods prices — the latter seems to be more critical for achieving superior out-of-sample forecasting performance and good in-sample fit.

6 Extending the Model to the Bretton-Woods Period

Given the favorable results of the model so far, my final step is to investigate whether the model also works on an extended data set. While a test of an empirical model on any data that has not been used in deriving the model is in itself of interest as it conveys important information on the model's stability, in the context of an empirical exchange rate model the natural challenge is to extend the model to the time before the break-up of the Bretton-Woods system of fixed exchange rates. The challenge is to explain the stylized fact that with the move from fixed to floating exchange rates the volatility of nominal and real exchange rates increased dramatically (Mussa 1986, Obstfeld 1995). Since it has proven to be extraordinarily difficult to identify only a single other macroeconomic variable with a similar regime-dependent variability (Baxter and Stockman 1989, Flood and Rose 1995), Flood and Rose (1995) proposed to use the ability to account for the fact that exchange rate volatility is substantially higher in floating rate regimes than during regimes of fixed rates as a simple "specification test" for empirical exchange rate models.

Theoretically, the hypothesis that the reduced form relationship between the real interest rate differential and the real exchange rate remains unaffected by the exchange rate regime is not implausible. It is governed by one parameter, Q_k , only, and this parameter - which determines the speed with which goods prices adjust - should not be affected strongly by a change in the exchange rate system. Thus, if uncovered interest parity held under fixed exchange rates as it did under floating rates and Q_k did not change, the parameter of the real interest differential in equation (10) should have been unaffected by a regime change if the choice of model has otherwise been correct.

Equation (12) presents the results from estimating the model with data running from 1963:1 through 1998:3.³⁰ To model the regime change in 1973:2 and the revaluation of the mark against the dollar in 1969:3 two step dummies which are 0 before the respective date and 1 thereafter were added to the model. A third step dummy was found significant in 1965:3 and also left in the model although no clear economic interpretation could be found for it. The step dummies are presented within the long-run relationship in squared brackets to facilitate the measurement of their impact on the intercept.

³⁰ Since the Bundesbank only started recording data on German net foreign assets in the fourth quarter of 1971, all entries before that date were set equal to the value of 1971:4 implying that the net foreign asset position did not change from 1963:1 to 1971:3. While this is certainly not the case, the error made by this assumption seems to be tolerable since trade balance data shows that current account imbalances between the United States and Germany have probably been quite small during this period.

$$\begin{aligned} \Delta q_t^T = & -0.308 \left[4.933 + 3.799(r^T - r^{T*})_{t-1} - 0.187f_{t-1}^c \right. \\ & (9.3) \quad (417.7) \quad (13.4) \quad (10.2) \\ & \left. + 0.105(S65:1 - S69:3 - S73:2) \right] \\ & (5.9) \\ & + 0.151(\Delta q_{t-1}^T + \Delta q_{t-3}^T + \Delta q_{t-4}^T) + (0.308 \cdot 3.799)\Delta(r^T - r^{T*})_t + 0.813\Delta(r^T - r^{T*})_{t-4} \\ & (3.3) \quad (8.9) \quad (4.4) \\ & + (0.309 \cdot 0.187)(\Delta f_{t-1}^c + \Delta f_{t-2}^c) + 0.005\Delta f_{t-3}^S + 0.014\Delta f_{t-4}^S \\ & (6.5) \quad (1.6) \quad (4.2) \\ & + 0.093(I74:1 + I74:3) \\ & (4.2) \end{aligned} \tag{12}$$

R^2 -ECM = 0.55; R^2 -Bewley = 0.74; $s = 0.030$, $T = 143$ [1963:1-1998:3],
 DW : 2.03, AR1: 0.825, AR1-4: 0.754, AR1-8: 0.378,
 BJ (Normality): 0.368, ARCH(4): 0.397, LM-Het: 0.182, RESET: 0.475
 ADF:-5.35 [3.74], x_m : 84.19 [15.22]
 $C_{exo}^2(2)$: 3.142 [5.99], $t_{r^T - r^{T*}}$: 0.022, t_{f^c} : -0.159

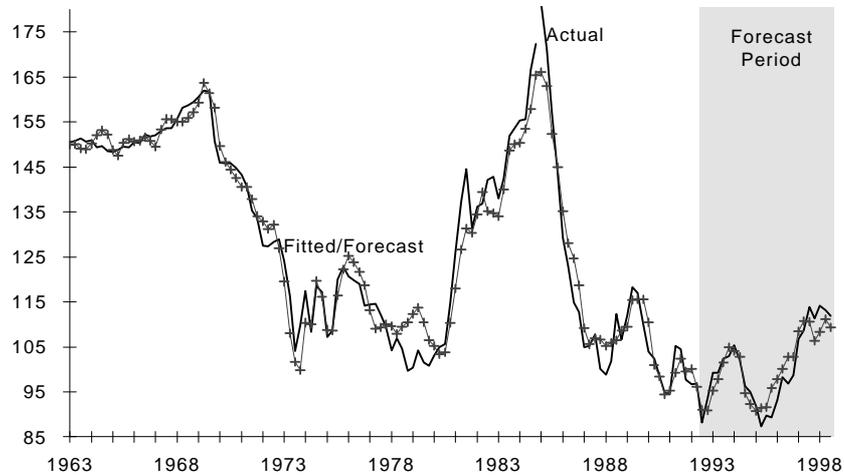
Remarkably, although more than ten years of data which stem from a different institutional setting have been added to the sample, both the short-run dynamics and the long-run coefficients are largely identical with those in equation (10). Again, the various tests do not indicate any misspecification. The fit of the ECM is nearly as good as before, that of the Bewley-transformed regression even increases to 74 %; the difference is due to the additional step dummies. The statistics of the cointegration tests are also higher in absolute terms than for equation (10).³¹ Weak exogeneity of the conditioning variables is, in contrast to the post-Bretton-Woods-sample, not rejected here. Moreover, a standard Chow-test rejects the hypothesis of a structural break in 1973:1 with $F_{10,122} = 4.284$. Given the small differences in the coefficient estimates and the good fit of the equation (see also Figure 5), it is not surprising that the predictive performance of equation (12) is also similar to that of equation (10). I estimated (12) from 1963:1 through 1992:2 and, again, used the last 25 quarters for the forecast exercise described above. Again, the random walk forecasts are by far outperformed, with the RMSE of the model being at their minimum a quarter of that of the naive model.³²

As the functional relationship between the real exchange rate and the fundamentals has apparently not been affected by the change in the exchange rate system the volatility

³¹ For the ADF test, the step dummies S65:1, S69:3 and S73:2 were included in the long-run relationship tested for non-stationarity.

³² The U-ratios for forecast horizons of 1,2,3,4,8 and 12 quarters, are 0.67, 0.54, 0.42, 0.33, 0.24, 0.27, respectively.

Figure 5:
In-Sample Fit and Out-of-Sample Forecasts of the
Conditional ECM (1963:1-1992:2)



Note: Quarterly data, 1963:1-1998:3. Forecasts based on equation (12) estimated through 1992:2. Out-of-sample forecasts generated from 1992:3 onwards.

of at least one of the two macroeconomic variables in this relationship must also have increased with the move to flexible exchange rates. Obviously, one sure candidate for a higher volatility under floating exchange rates is the net foreign asset variable as it is common knowledge that the United States only started to become a significant debtor country in the early 1980s. Unfortunately, I cannot give any formal statistical support for this hypothesis due to the lack of hard data on foreign assets in the 1960s. For the real interest rate differential, casual observation of Figure 1 suggests that its volatility may also have been higher under floating rates, and this impression can be tested formally.

Table 5 presents the means and standard deviations of the real exchange rate and the differences of real and nominal interest rates for the period of 1961:3 to 1971:4 on the one hand and the floating rates period, starting in 1973:2, on the other.³³ To control for changes in means between the two periods, the analysis of the series' volatility focuses on the coefficients of variation which normalize the standard deviations by the respective means. The first column presents the data for the real exchange rate based on tradables prices q_t^T , reproducing the well-known result of a dramatic increase in volatility with the move to floating nominal exchange rates; its coefficient of variation for the floating rates period is 4.3 times that of the fixed rates period. The second column shows the results for the real interest rate differential $r_t^T - r_t^{T*}$; its coefficient of variation has increased by a factor of 3.0 with the move to flexible rates. By implication, this variable

³³ The period of 1972:1 to 1973:1 has been omitted as it "formally" still belongs to Bretton-Woods period but is already characterized by high real exchange rate volatility as the fixed rate system was already under strain.

Table 5:
Variability of the real exchange rate and real interest rates
under fixed and under floating nominal exchange rates

	q_t^T		$r_t^T - r_t^{T*}$		$r_t - r_t^*$	
	61-71	73-98	61-71	73-98	61-71	73-98
Mean	5.015	4.728	-0.011	-0.010	-0.009	-0.013
Std. Dev.	0.039	0.161	0.015	0.040	0.019	0.024
$cv = \frac{ \text{Std. Dev.} }{\text{Mean}}$	0.008	0.034	1.335	4.027	2.079	1.809
cv_{73-98}/cv_{61-71}	—	4.3	—	3.0	—	0.9

Notes: Real exchange rate as defined above measured as an index (1990=100), real interest rate differential measured in percentage points. Precise sampling periods are 1961:3 - 1971:4 and 1973:2 - 1998:3.

alone is capable of explaining nearly three quarters of the increase in the volatility of the real exchange rate. Add to this the higher variability of the current account in the floating rates period, especially since the early 1980s, and most of the real exchange rates regime-dependent volatility seems to be explainable by the volatility of fundamentals. By implication the model passes the specification test proposed by Flood and Rose (1995) in that it is capable of explaining the regime dependence of the volatility of the real exchange rate: the increase in the volatility of the real exchange rate with the move from fixed to floating exchange rates coincides with an increase of the variability of the (tradables inflation-based) real interest rate differential and possibly also that of the net foreign asset position of the United States vis-à-vis Germany.

Regime-dependent volatility can, however, not be found for the conventionally defined real interest differential $r_t - r_t^*$, the variability of which even seems to have declined somewhat after the break-up of the Bretton-Woods system; a researcher looking for a model specification that passes the Flood and Rose (1995) test would thus reject a specification based on the conventional real interest rate differential. Again, the explicit consideration of the Balassa-Samuelson effect shows to be of particular importance for achieving a stable model. The finding of similar volatility patterns between real exchange rates and real interest rates in this study may, moreover, have been enhanced by the use of quarterly instead of monthly data as quarterly averaging probably reduces the volatility of the post-Bretton-Woods real exchange rate without having much effect on real interest rates. This is another argument in favor of using data of lower frequency when sorting out the impact of fundamentals on exchange rates.

7 Concluding Remarks

This paper has shown that more than half of the quarterly variability of the dollar's exchange rate vis-à-vis the D-mark can be explained by a small set of macroeconomic fundamentals comprising short-run nominal interest rates, producer prices in manufacturing and bilateral net foreign asset holdings. The findings are consistent with theoretical exchange rate models based on the assumption that goods prices are sticky in the short run. They also seem to support the notion of short-run exchange rate overshooting. Apart from that, however, they highlight the importance of accounting for the effects of non-traded goods and current account imbalances on the real exchange rate.

As I have confined my interest to the D-Mark/U.S. dollar parity so far, a natural extension of this work would be to transfer the model to another pair of currencies. This may, however, not be a simple task. One problem is that the measure for the stock of net foreign assets used in the present study is not available for use with other currencies vis-à-vis the U.S. dollar. Still, the dollar's behavior is definitely the most challenging for an empirical exchange rate model to explain.³⁴ The data problem apart, it is the fact that all of the main more or less freely floating currencies seem to have their idiosyncrasies which caution against a simple transfer of the present model. The Japanese yen, which would be an interesting candidate to analyze, seems to have been subject to official exchange rate management over long parts of the floating rates period with frequent foreign exchange interventions and capital and exchange controls (Henning 1994, p. 121ff.). For the British Pound, the oil discoveries in the 1970s may have interfered with its real exchange rate and it also makes the choice of a price index for which PPP may hold more difficult. In any case, a modeling strategy that is to produce good out-of-sample forecasts will probably have to take account of these features. I leave this for future work to attempt.

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³⁴ See also Lothian (1998) on the apparently rather special behavior of the U.S. dollar during the floating rates period.

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