How stable are exchange rate determination equations of asset approaches? A time-varying coefficient approach

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February 2009

Abstract

This paper examines the importance of different fundamental regimes by applying asset market models on exchange rates to the most important exchange rate, namely the Deutsche Mark (Euro)-US Dollar exchange rate. We test for different hypotheses: firstly, there is no stable long-run equilibrium relationship among fundamentals and exchange rates since the breakdown of Bretton Woods. Secondly, there are no perseverative regimes, i.e. either the coefficient values for the same fundamentals differ or the significance differs. Thirdly, there is no regime in which no fundamentals enter. Fourthly, the step-wise relationship acts as an error correction term.

JEL Classification: E44, F31, G12

Keywords: Structural exchange rate models, cointegration, structural breaks

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1. Introduction

Exchange rate economics is still one of the most controversial research areas in economics. After the first generation models of exchange rate determination which see the exchange rate as the relative price of domestic and foreign monies (e.g. Dornbusch, 1976; Frenkel, 1976; Kouri, 1976; Mussa, 1976) was brought to the data, it became clear that exchange rate models can only partly be used to explain historical exchange rates with the help of fundamentals and perform poorly in forecasting, in particular (Meese and Rogoff, 1983). The results of the seminal paper by Meese and Rogoff (1983) still represent the benchmark: exchange rate forecasts by structural models can hardly outperform naïve random walk forecasts. Since the publication of the Meese and Rogoff paper many contributions have tried to refute their results. By using the implicit assumption that exchange rates and fundamentals are cointegrated and by implementing exogenous parameter restrictions a couple of authors can find predictability in the long run for a similar period as in Meese and Rogoff (e.g. Mark, 1995; Chinn and Meese, 1995).[3] However, an extension of the sample yields mostly contrary findings (e.g. Kilian, 1999; Abhyankar, Sarno and Valente, 2005). A critical point is the implicit assumption of cointegration which is inadequate because it leads to biased conclusions (Berkowitz and Giorgianni, 2001).

While the models of the late 1980s mostly neglect a long-run relationship between the fundamentals and the exchange rate, structural models which test explicitly for a long-run relationship among exchange rates and fundamentals were applied at the beginning of the 1990s. These kinds of models which base upon cointegration techniques, can indeed improve the evidence in favour of predictability in the long-run when periods up to the end of the 1990s are discovered (MacDonald and Taylor, 1993, 1994).[4] However, an extension of the period of observation yields a breakdown in cointegration relationships (e.g. Groen, 1999). While investigations directed solely to bilateral exchange rates mostly cannot find both long-run relationships among exchange rates and monetary fundamentals and predictability, panel estimation exercises give evidence in favour of both (Groen, 2000, 2002; Mark and Sul, 2001; Rapach and Wohar, 2004). Surprisingly, less attention is directed to a closer examination of the link between exchange rates and fundamentals with respect to structural changes if cointegration does not hold.

Stock and Watson (1996) show that univariate and bivariate macroeconomic time series are subject to substantial instabilities which result in poor forecasting performances. Goldberg and Frydman (1996a, 2001) provide evidence that periods exist in which the monetary model

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[3] Mark (1995) is the first author who focuses on more exchange rates simultaneously. He includes the Canadian Dollar, the Deutsche Mark, Japanese Yen and Swiss Franc expressed in US Dollar. In Chinn and Meese (1995) the Swiss Franc is not included but the Pound Sterling in US Dollar as well as the US Dollar and the Deutsche Mark in Japanese Yen.

[4] MacDonald and Taylor (1994) investigate the Pound Sterling-US Dollar exchange rate.

is valid and in which it is not. Thus, the instability of the monetary model in reality can be an explanation for the findings of Cheung et al. (2005). In the recent past, models are applied to the monetary approach which are able to take account of different regimes. Sarno, Valente and Wohar (2004) use a Markov regime switching model in order to investigate the response of exchange rates on deviations from fundamental values in different regimes. De Grauwe and Vansteenkiste (2007) investigate particularly the adjustment under different inflation regimes. Taylor and Peel (2000), Taylor, Peel and Sarno (2001) and Kilian and Taylor (2003) make use of models that allow for smooth transition between two states supporting the hypothesis that the link between exchange rate adjustment towards equilibrium paths is nonlinear, concretely, fundamentals become important if the deviation from an equilibrium rate is large. Frömmel, MacDonald and Menkhoff (2005a,b) test directly for different regimes in the exchange rate determination equation of the real interest rate differential model. Nevertheless, the latter are the only one who allowed the coefficients in the exchange rate determination process itself to change. Since Frömmel et al. (2005a,b) specify their model in first differences a true long-run relationships is not investigated.[5] All other contributions direct their attention on deviations from a fundamental value which assumes cointegration with implied restrictions without modelling the long-run structure separately.

However, both mentioned regime switching approaches have in common that they only allow for a fixed number of regimes, whereas the regimes are perseverative. In early works, Schinasi and Swamy (1989) and Wolff (1987) apply a time varying coefficient model (TVP) on monetary models. They show that their models are more adequate in forecasting than fixed coefficient models. Obviously, time-varying parameters are very important.

The results of market surveys exhibit that foreign exchange traders report that different fundamentals are important during different periods (e.g. Gehrig and Menkhoff, 2006). From imperfect knowledge theory it can be derived that combinations between different fundamentals must not be systematically similar (e.g. Frydman and Goldberg, 1996b). Hence, it is reasonable to assume that a strong relationship between exchange rates and fundamentals exist during sub-periods but its nature will change considerably over time. From this point of view, a fundamental value of the exchange rate will exist in the sense that a part of the exchange rate is driven by fundamentals. For this reason, a positive analysis should be applied instead of a normative one.

Our aim is to test for different hypotheses: firstly, there is no stable long-run equilibrium relationship among fundamentals and exchange rates since the breakdown of Bretton Woods. Secondly, there are no perseverative regimes, i.e. either the coefficient values for the same fundamentals differ or the significance differs. Thirdly, there is no regime in which no fundamentals enter. Fourthly, the step-wise relationship acts as an error correction term.

[5] In order to obtain a long-run perspective, they use annual changes on a monthly data set.

The rest of this paper is organized as follows. We start with a multiple structural change model developed by Bai and Perron (1998, 2003) which is applied to the reduced form of structural exchange rate models. Hypothesis 1 is accepted if more than 1 structural change is found. In the next section, we use the estimated break points to generate indicator functions with which help we estimate the structural model in order to obtain estimates for the different regimes. For this purpose, we apply the fully modified OLS estimator by Phillips and Hansen (1990), which is able to handle with non-stationary variables as regressors and regressands. The results are then evaluated with respect to the second and third hypothesis. Finally, we construct an error correction term and regress the change of the exchange rate on this error correction term to investigate whether the exchange rate adjusts.

2. Exchange rate modelling

2.1 Economic Theory

After the breakdown of Bretton Woods at the beginning of the 1970s exchange rate models were developed which see exchange rates as asset prices (e.g. Dornbusch, 1976a; Frenkel, 1976; Kouri, 1976). The resulting class of monetary models of exchange rate determination rests on the assumption that the exchange rate equilibrates the demand for and the supply of the stock of domestic and foreign assets. Thus, the exchange rate results from the agents' willingness to hold domestic and foreign assets (e.g. Mussa, 1976).

All models of this kind have in common that they rely on the money demand function of the form

$$\frac{M}{p} = L(Y^r, i) \qquad (1)$$

with M the money supply, p the price level, L the money demand which depends on real income (Y^r), and interest rates (i). A basic assumption of the standard monetary model is that the purchasing power parity (PPP) holds. In the log-linearized form, the exchange rate can be expressed as the differences in price levels which via money market equations is equal to the difference between domestic and foreign money supply less real money demand, so that the exchange rate determination

$$\begin{aligned} s &= \alpha + (\beta_1 m - \beta_2 y + \beta_3 i) - (\beta_1^f m^f - \beta_2^f y^f + \beta_3^f i^f) \\ &= \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f + \beta_3 i - \beta_3^f i^f \end{aligned} \qquad (2)$$

results. In the literature this model is widely known as the Frenkel-Bilson (FB) model.[6] In the original monetary model α is zero and $\beta_1 = \beta_1^f = 1$ because of the structure of the money

[6] β_j are elasticities and α is a constant term. m and y are the logarithms of money supply and real income. The interest rates are expressed as percentage.

demand function. Equation (2) can be rewritten with the restriction that the (semi-) elasticities of the interest rates are equal. Thus

$$s = \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f + \beta_3 (i - i^f). \qquad (3)$$

If the uncovered interest rate parity (UIP) holds, $(i - i^f)$ can be replaced by the expected change in the exchange rate $(E_t(s_{t+1}) - s_t)$. With an expectation generating mechanism based upon PPP the differences in interest rates can then be replaced by the differences in expected rates of inflation.[7] Since it is known that the exchange rate often deviates from the PPP the adjustment towards the PPP value can be taken into account in addition to the expectations concerning the expected rates of inflation $E_t(s_{t+1} - s_t) = -\phi(s_t - \bar{s}) + \pi_t - \pi_t^f$.[8] The real interest rate model (RID) by Frankel (1979) arises if the expectation formation process is combined with the UIP and solved for the expected change in the exchange rate (equation (4)).

$$s = \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f - \beta_3 (i_t - i_t^f) + \beta_4 (\pi_t - \pi_t^f). \qquad (4)$$

The exchange rate decreases if a positive interest rate differential exists and increases if a positive inflation rate differential prevails. With the help of equation (4) a similar process can be explained as in the overshooting case of Dornbusch (1976a). In Dornbusch (1976a) the exchange rate is negatively correlated with the interest rate differential but without a feedback on inflation expectations, i.e. β_4 is zero. Equation (4) allows the exchange rate to deviate from the PPP in the short-run, i.e. it reacts negatively on interest rates, but still positively on inflation rate expectations. Following Frankel (1979) exactly, β_1 and β_1^f must be equal to one. [9]Since a distinction must be made between the Dornbusch model and the Frankel model we refer to the RID model when talking about equation (4).

A weakness of the traditional monetary model is that the real exchange rate must be constant in the long-run. In an investigation which examines the exchange rate behaviour over a long period it is reasonable to assume that real shocks change the real exchange rate which has consequences on the nominal exchange rate. In order to take account of real shocks, Hooper and Morton (1982) introduce changes of the equilibrium real exchange rate into the traditional monetary model. In addition to nominal impact factors, the real side of the economy is introduced by taking account of innovations in the current account. The equilibrium of the real exchange rate depends on the desire of domestic and foreign agents to accumulate (or decumulate) net foreign assets in the long run. Since the desire of

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[7] This formulation is equivalent to a money demand function in which the expected rates of inflation enter as opportunity costs.

[8] ϕ denotes the adjustment speed to the equilibrium value $\bar{s}$. π is the expected rate of inflation.

[9] Nevertheless, Driskill und Sheffrin (1981) show that overshooting requires $\beta_1 > 0$ and $\beta_1^f < 0$.

accumulating (or decumulating) net foreign assets is reflected by the equilibrium current account surplus, the equilibrium real exchange rate is linked to the equilibrium net foreign asset position and the equilibrium current account position. An unexpected rise in the current account means that too much net foreign assets are accumulated which in turn reduces the demand for foreign capital and causes the domestic currency to appreciate nominally. Thus, unexpected (positive) shocks to the equilibrium net foreign asset position result in a nominal appreciation. Hooper and Morton (1982) proxy the net foreign assets by the cumulated current account and equation (4) can be extended by the cumulated trade balances as a proxy for the current account balance (equation (5)).[10]

$$s = \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f - \beta_3 (i_t - i_t^f) + \beta_4 (\pi_t - \pi_t^f) - \beta_5 CTB_t + \beta_5^f CTB^f . \quad (5)$$

The Hooper and Morton model is usually applied by estimating equation (5) with cumulated overall domestic and foreign trade balance. Without a loss in generality the cumulated overall trade balances can be replaced by the trade balances with the same meaning because the equilibrium change in the net foreign asset position is the equilibrium trade balance. In addition to the real exchange rate motive, Hooper and Morton (1982) also use the overall trade balances as an indicator for the risk premium which arise from government debt, an insufficient holding of international reserve and foreign indebtness. A fall in the net foreign asset position (in particular if it is negatively) increases the risk premium from which an increase in the exchange rate follows. Hence, the risk premium will sensitively react to a worsening of a negative net foreign asset position. In a bilateral case it is straightforward to use the bilateral cumulated trade balance (BCTB) instead of the overall cumulated trade balances (equation (6)).

$$s = \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f - \beta_3 (i_t - i_t^f) + \beta_4 (\pi_t - \pi_t^f) - \beta_5 BCTB_t . \quad (6)$$ [11]

Since it is expected that the PPP holds for traded goods rather than for a mixture of traded and non-traded goods as implicitly assumed by using the overall price index, the prices of traded goods can be taken into account (e.g. Dornbusch, 1976b). If the overall price index, which is determined by the money market, consists of prices of both traded and non-traded goods and if the PPP is only valid for traded goods, the monetary approach yields an exchange rate determination equation of the form

$$s = \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f - \beta_3 (i_t - i_t^f) + \beta_4 (\pi_t - \pi_t^f) + \beta_6 \frac{P_t^T}{P_t^{NT}} - \beta_6^f \frac{P_t^{fT}}{P_t^{fNT}} . \quad (7)$$ [12]

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[10] Since data on the current account are not available at a monthly frequency, it is adequate to proxy the current account by the trade balance.

[11] A caveat by using the cumulated bilateral trade balance as a proxy for net foreign assets is that only a part of the current account is covered. Besides the transfers, income and trade in services is excluded. Since in income returns of capital dominate it depends predominantly on return such as interest rates which are included. Since trade in services is a minor issue it is reasonable to exclude it.

[12] T denotes trabables and NT non tradables.

The proportion of traded to non-traded goods reflects the real exchange rate. An increase in the price of tradables relative to the price of non-tradables will cause the nominal exchange rate to increase because the domestic good is substituted by the foreign good. In the flex price model β_4 is equal to zero and the exchange rate reacts positively on the interest rate differential (Wolff, 1987).

In applied monetary models equation (2) is typically estimated based upon a reduced form in which it is assumed that the elasticities for an economic variable are identical in both countries $\beta_1 = \beta_1^f$, $\beta_2 = \beta_2^f$ and $\beta_3 = \beta_3^f$ (e.g. Meese and Rogoff, 1983). An analysis in which the coefficients are restricted to be equal for each variable will typically result in biased coefficients (e.g. Haynes and Stone, 1981). If the structure of the economy is not known a priori, restricted coefficients do not help in explaining the exchange rate. While the traditional monetary model assumes that domestic and foreign assets are perfect substitutes the assumption is relaxed by highlighting the role of risk as described by Hooper and Morton. A model that takes explicitly account of risk premiums is the portfolio balance models (Branson, 1977). If a risk premium becomes more important, it is preferable to use the portfolio balance approach. In the following we make use of a hybrid model which catches effects that can be found both in monetary and portfolio models (e.g. Frankel, 1983). As a consequence, we remove the restrictions of equal parameters of the interest rate differential and the inflation rate differential in equations (4), (5),(6), and (7).

Thus, we start as less restrictive as possible and we bear in mind dynamics stemming both from the portfolio balance approach and the monetary approach. Finally, we have three different models which all rely on the baseline specification of the unrestricted RID model exchange rate determination equation in equation (4). The models are the (unrestricted) traditional RID model and two extensions: one with the cumulated overall trade balances (equation (5)) and one with the bilateral cumulated trade balance (equation (6)).

2.2 Long-run analysis with time-varying coefficients

Explicitly, we assume that the structure of the economy is not known, i.e. all coefficients are unrestricted. Wolff (1987) mentions three reasons why a time-varying coefficient model should be superior to fix coefficient models. First of all, the money demand function is subject to instabilities which cause the coefficients in the exchange rate determination equation of a reduced model to change. Another reason is the famous Lucas critique: coefficients change if a change in the policy regime occurs. A last point is directed to the long-run real exchange rate. The monetary model assumes that the purchasing parity holds in the long-run from which follows that the long-run real exchange rate is stable. Innovations to the real exchange rate from the real side of the economy can lead to changes in the coefficients. Only the last

issue deserves less attention in our analysis with respect to the choice of the estimation technique because we explicitly account for changes in the real exchange rate.
A motivation for time-varying coefficient models can also be derived from different theories. In intertemporal open macroeconomic models (e.g. Obstfeld and Rogoff, 1995), money does not depend on income, here it depends on real consumption. If we proxy real consumption by real income, a change in the average rate of consumption will result in a change in the elasticity of income in the exchange rate equation. Thus, if consumption shares vary, which is, for instance, true for the USA, the exchange rate determination equation will be time-varying.
As Wilson (1979) argues an anticipated policy change, i.e. an expansionary monetary policy, can generate dynamics which are different from unanticipated changes. In Wilson (1979) the overshooting dynamics are slightly different from those in Dornbusch (1976a). A very important result is that an appreciation period of the domestic currency coincides with the increase in money supply while in the Dornbusch model a boost in money supply coincides with a depreciation. If anticipated and unanticipated shocks alternate, fixed coefficient models are inadequate because they cannot catch both effects simultaneously.
As shown by Sarno, Valente and Wohar (2004) or De Grauwe and Vansteenkiste (2007), the adjustment of the exchange rate towards the long-run equilibrium relationship seems to be not time-invariant. Over a long span of data, we expect that adjustments differ from period to period. An adjustment towards the long-run equilibrium relationship can occur because the exchange rate reacts predominantly on the fundamentals or the fundamentals react to changes in exchange rates. In the last case, it is possible that the exchange rate does not adjust in sub-periods. Consequently, the adjustment coefficient can differ.
A concept in which these issues can be built is outlined by Siklos and Granger (1997). They point out that a cointegration relationship can be subject to structural changes and argue that the common stochastic trends are only present in specific periods. In this respect they introduce the concept of regime-sensitive cointegration, or “switch on – switch off” cointegration. Consequently, the concept of regime-sensitive cointegration can be combined with a time-varying coefficient approach.
Let X_t^1, X_t^2 and Y_t be different processes where

$$Y_t = \beta_t^1 S_t^1 + \beta_t^2 S_t^2 + \phi_{1t} Z_t + \varepsilon_t^y \quad (8)$$

$$X_t^1 = S_t^1 + \phi_{2t}^1 Z_t + \varepsilon_t^{x1} \quad (9)$$

$$X_t^2 = S_t^2 + \phi_{2t}^2 Z_t + \varepsilon_t^{x2} \quad (10)$$

S_t and Z_t are both I(1) but do not share a common stochastic trend. ε_t^x and ε_t^y are both i.i.d. error processes which follow a normal distribution with zero mean. Furthermore, β_t^k can be a time-varying cointegration parameter, i.e.

$$\beta_t^k = 1_{1t}^k \beta_t^k + ... + 1_{mt}^k \beta_t^k \quad \text{with } k = [1,2] \tag{11}$$

with

$$1_{mt}^k = 1(T_{j-1} < t < T_j), \text{ with } j = 1,...,m. \tag{12}$$

In equation (12) it is not allowed that the time periods overlap so that the cointegration parameter is permitted to be absent during sub-periods. From this follows that one of the two common stochastic trends can vanish in equation (8).

Cointegration of X_t^1, X_t^2 and Y_t requires that a linear combination of X_t^1, X_t^2 and Y_t with cointegration vector of $(1, \beta^1, \beta^2)'$ is stationary. Hence, the linear combination is

$$\begin{aligned} Y_t - \beta_t^1 X_t^1 - \beta_t^2 X_t^2 &= \beta_t^1 S_t^1 + \phi_{1t} Z_t + \varepsilon_t^y - \beta_t^1 S_t^1 - \phi_{2t}^1 Z_t - \beta_t^1 \varepsilon_t^{x1} - \beta_t^2 S_t^2 - \phi_{2t}^2 Z_t - \beta_t^2 \varepsilon_t^{x2} \\ &= (\phi_{1t} - \beta_t^1 \phi_{2t}^1 - \beta_t^2 \phi_{2t}^2) Z_t + \varepsilon_t^y - \beta_t^1 \varepsilon_t^{x1} - \beta_t^2 \varepsilon_t^{x2}. \end{aligned} \tag{13}$$

Equation (13) is a cointegration relationship if $\phi_{1t} - \beta_t^1 \phi_{2t}^1 - \beta_t^2 \phi_{2t}^2$ is zero and the stochastic trend Z_t vanishes so that cointegration is switched on. Similarly to equations (11) and (12), a time varying representation of ϕ_t and ϕ_t^k can be achieved. For this reason, they depend on time and can be present or absent in sub-periods. This result is independent from the number of common stochastic trends involved in the system. If the condition is not valid, cointegration is switched off. The combination of equations (8), (9), (10), (11), and (12) shows that the system is driven by two common stochastic trends which can be absent in subsequent periods. If a system has at least one continuous common stochastic trend, Y_t will only continuously cointegrate with X_t^k under the condition that $\phi_{1t} - \beta_t^1 \phi_{2t}^1 - \beta_t^2 \phi_{2t}^2$ is zero. The error correction term is therefore with $ect_t = \varepsilon_t^y - \beta_t^1 \varepsilon_t^{x1} - \beta_t^2 \varepsilon_t^{x2}$

$$ect_t = Y_t - \beta_t^1 X_t^1 - \beta_t^2 X_t^2 \tag{14}$$

from which the error correction form for Y_t follows

$$\Delta Y_t = -\alpha_{1t}(Y_t - \beta_t^1 X_t^1 - \beta_t^2 X_t^2) + \eta_t \tag{15}$$

in which η_t is a i.i.d. variable which follows a normal distribution with zero mean. If one of the stochastic trends in equation (8) is currently absent the corresponding X_t^k variable will not enter the cointegration vector and the cointegration vector will only contain two elements. It can be seen that cointegration is continuously present over the whole period of observation but only the composition of the cointegration vector changes. In addition to a time-varying cointegration vector, we assume that the causality between the variables can change during the period of observation. This means, that the dimensionality of the vector which contains the adjustment coefficients can be reduced during sub-periods. Assuming that the adjustment of the X_t^k is still present, as long as cointegration prevails, $-\alpha_{1t}$ in equation (15

) changes not only its magnitude, it can also be zero if Y_t does not adjust to the long-run relationship.

In a long-run relationship analysis we can be confronted with both switch on and off cointegration and a changing cointegration vector. Finally, our approach takes account of different regimes. It is able to distinguish between cases when the cointegration relationship is switched off or different adjustments are present.

For a multivariate case

$$Y_t = \beta_t X_t + \tau \qquad (16)$$

with

$$X_t = [X_t^1, ..., X_t^k] \text{ for } n = 1, ..., K \qquad (17)$$

with K as the maximum number of explanatory variables. The matrix X_t has the dimension $(K \times 1)$ and β_t the dimension $(1 \times K)$. For the empirical analysis, we consider the following models:

Model one:

$$Y_t = [s_t], \qquad X_t = \left[m, y, i, \pi, m^f, y^f, i^f, \pi^f,\right] \qquad (18)$$

Model two:

$$Y_t = [s_t], \qquad X_t = \left[m, y, i, \pi, m^f, y^f, i^f, \pi^f, CTB, CTB^f\right] \qquad (19)$$

Model three:

$$Y_t = [s_t], \qquad X_t = \left[m, y, i, \pi, m^f, y^f, i^f, \pi^f, BCTB\right] \qquad (20)$$

Model four:

$$Y_t = [s_t], \qquad X_t = \left[m, y, i, \pi, m^f, y^f, i^f, \pi^f, \frac{p^T}{p^{NT}}, \frac{p^{fT}}{p^{fNT}}\right] \qquad (21)$$

3. Modeling structural changes and estimating cointegrating relations : Methodological issues

3.1 Testing for multiple structural changes

In general, two frameworks for the testing of structural changes can be distinguished: Generalized fluctuation tests fit a model to the data and derive an empirical process that captures the fluctuations either in the residuals or in parameter estimates. If the generated process exceeds the boundaries of the limiting process which can be derived from the functional central limiting theorem the null of parameter constancy must be rejected, meaning a structural change occurs at the corresponding time (Zeiless et al, 2003). Well known examples of these methods are the classical and the OLS based cusum test and the fluctuation test of Nyblom (1989). Those structural change tests are predominantly designed for stationary variables. In the case of a cointegration analysis an Eigenvalue fluctuation test developed by Johansen and Hansen (1999) can be applied which bases upon Nyblom. While these procedures have the advantage of not assuming a particular pattern of deviation from the null hypothesis they can only identify a single break or show generally instability. The other way to test for structural changes is to compare the OLS-Residuals from regressions for different subsamples. This can for example be done by applying F-Statistics or the Chow test. In this paper, we will adopt an extension of the latter case developed by Bai and Perron (1998, 2003). Their basic idea is to choose such break points that the sum of squared residuals for all observations is minimized.

As a starting point, consider a multiple linear regression with m break points and m+1 regimes

$$y_t = x_t'\kappa + z_t'\delta_j + u_t \ , \qquad (t = T_{j-1}+1, \dots, T_j) \qquad (22)$$

for j=1,……m+1 with the convention that T_0=0 and $T_{m+1} = T$

y_t is the the dependent variable, x_t' and z_t' denominate the regressors and κ and δ are the coefficient vectors.

With a sample of T the first step is to calculate the corresponding values for all possible T(T+1)/2 segments.[13] The estimated break points $(T_1 \dots\dots T_m)$ by definition represent the linear combination of these segments which achieve a minimum of the sum of squared residuals (Bai and Perron, 2003). Formally

$$(\hat{T}_1, \dots, \hat{T}_m) = argmin_{T_1,\dots,T_m} S_T(T_1, \dots, T_m) \qquad (23)$$

Bai and Perron (2003) develop a dynamic programming algorithm which compares all possible combinations of the segments. Their methodology allows testing for multiple structural breaks under different conditions.[14] Within our framework, the location of the break points is also obtained by the sum of squared residuals. To select the dimension of the model we apply the Bayesian Information Criterium (BIC) which according to Bai and Perron (2001) works well in most cases when breaks are present. After calculating the tests for all

[13] Bai and Perron (1998) note that for practical purposes, less than T(T+1) segments are permissible, for example if a minimum distance between each break may be imposed. In the framework of this paper, breaks are allowed to occur every 12 months.

[14] One possibility is to test the null of no change against the hypothesis of a fixed number of breaks m=k using F tests based on the sum of squared residuals under both hypotheses. For an unknown number of breaks, one way is to allow a maximum number of breaks. In this case one can apply the so called double maximum test. The number of break points is then selected by comparing the F-values described above for the different numbers of break points and select the configuration with the highest F-value respectively the minimum of the sum of the squared residuals. Another possibility is to test sequentially for an additional break using the l vs. l+1 break tests. For Details see Bai and Perron (1998, 2003).

possible break points the sequence $(\hat{T}_1, \ldots, \hat{T}_m)$ is selected as the configuration at which the BIC achieves its minimum. Carrioni-Silvestre and Sanso (2006) show that this approach leads to a consistent estimation of the break fraction. Note that the break points obtained in this fashion are a local minimum of the sum of squared residuals given the number of break points but not necessary a global minimum.

It is important to note that the procedure of Bai and Perron was originally developed for the case of stationary variables. Nevertheless, it can also be applied to I(1) variables. Siklos and Granger (1997) use this methodology to identify structural breaks in the Interest Parity between the United States and Canada in the context of regime-sensitive cointegration. Zumaquero and Urrea (2002) point out that the break estimator is consistent in the non-stationary case. Using disaggregated price indexes for seven countries they test for structural breaks in the coefficients of cointegrating relations representing absolute and relative Purchasing Power Parity. They also examine instabilities in the adjustment behavior of price rations and exchange rates. Finally, Perron and Kejriwal (2008) showed that the results of Bai and Perron (1998) in general continue to hold even with I(0) and I(1) variables in the regression.[15] This is also true if one allows for endogenous I(1) regressors.[16] The use of information criteria as the BIC is also allowed in both cases.

3.2 Estimating cointegrating relations with single equations

After identifying the break points we now turn to the problem of estimation. As Bai and Perrons methodology is designed for single equations, we cannot consider multivariate system estimators as proposed by Johansen (1988) or Stock and Watson (1988). Besides the traditional approach of Engle and Granger (1988), several modified single estimators have been developed. Examples are the fully modified estimator by Phillips and Hansen (1991) and the approach of Engle and Yoo (1991). For a review of the different estimation methods for cointegrating relationships see Hargreaves (1994), Phillips and Loretan (1991) and Capporale and Pittis (1999). Even in the case of a multi dimensional cointegration space, single equation approaches can be used to achieve asymptotically efficient estimates of single cointegrating relationships. However, with no long-run equilibria, i.e. with a rank of zero, one would obtain a spurious regression (Caporale and Pittis, 1999).

For our purposes, the fully modified estimator is the most suitable method. In contrast to traditional single equation formulars it considers endogeneous regressors (Phillips, 1991). Phillips and Hansen (1990) show that the FM-OLS estimator is hyperconsistent for a unit root in single equations autoregression. Phillips (1995) proves that the FM OLS procedure is

[15] This is only true if, as in this paper, the intercept is allowed to change across segments

[16] For the case without unit roots, Perron and Yamamoto (2008) show that in the presence of endogenous regressors the estimation of the break dates via OLS is preferable to an IV procedure.

reliable in the case of full rank or cointegrated I(1) regressors[17] as well as with I(0) regressors. Hargreaves (1994) runs a Monte Carlo Simulation and points out that single estimators in general are robust if more than one cointegrating relation exists with the FM OLS estimator doing best. He concludes that the FM OLS estimator should be preferred, even in advance to multivariate methods, if one wants to examine one cointegrating vector and is unsure about the cointegrating dimensionality. This is of particular interest for the aim of this paper as we are primarily interested in the long-run relationship between exchange rates and fundamentals and do not want to pay too much regard to other cointegrating relationships which might arise between the reported fundamentals. Caporale and Pittis (1999) claim that the FM-OLS estimator and the Johansen estimator perform best in finite samples. Furthermore, also Phillips and Hansen (1990), Hargreaves (1994) and Cappucio and Lubian (2001) report good finite samples properties of the FM OLS estimator.

The root idea of this concept is to estimate cointegrating relations directly by correcting traditional OLS with regard to endogenity and serial correlation (Phillips, 1995). Let z_t denominate an n-vector where y_t denotes an r dimensional I(1) process while x_t is an n-r =($(n-r)_1 + (n-r)_{2)}$ dimensional vector of cointegrated or possibly stationary regressors. u_t represents an n-vector stationary time series. Both vectors can be partitioned as followed.

$$z_t = \begin{bmatrix} y_t \\ x_{1t} \\ x_{2t} \end{bmatrix} \qquad u_t = \begin{bmatrix} u_{1t} \\ u_{2t} \\ u_{3t} \end{bmatrix}$$

The data generating mechanism for y_t is represented by the following cointegrated relation

$$y_t = \beta x_{1t} + u_{1t} \tag{24}$$

The vectors of the regressors are specified as follows

$$\Delta x_{1t} = u_{2t}$$

$$x_{2t} = u_{3t}$$

The estimator corrections can be applied without pre-testing the regressors for unit roots as both corrections can be conducted by treating all components of x_t as non stationary. For the nonstationary components, this transformation reduces asymptotically to the ideal correction while the differenced stationary components vanish asymptotically. Such a correction will have no effects for subvectors of x_t where serial correlation or endogeneity are not present.[18] A further advantage is that we do not have to account for cointegration between the x_{1t} regresors within this methodology (Phillips, 1995).

To imply the corrections, we first consider the long run covariance matrix Ω which can be decomposed into a contemporaneous variance and the sums of autocovariances (Hargreaves, 1994).

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[17] Note that the direction of cointegration needs not to be known. Regressors containing a deterministic trend are also allowed.

[18] Without serial correlation or endogeneity the FMOLS estimator is identical to the OLS estimator.

$$\Omega = \mathrm{E}(u_t u_t') + \sum_{k=2}^{\infty} \mathrm{E}(u_t u_k') + \sum_{k=2}^{\infty} \mathrm{E}(u_k u_t') \tag{25}$$

$$\Omega \ = \Sigma + \lambda + \lambda'$$

We define Δ as

$$\Delta = \Sigma + \lambda. \tag{26}$$

Estimation of these covariance parameters can be achieved by using the pre-whitened kernel estimator suggested by Andrews and Monahan (1992).[19] The endogeneity correction than has the form

$$y_t^* = y_t - \widehat{\Omega}_{0x} \widehat{\Omega}_{xx}^{-1} \hat{u}_{0t}. \tag{27}$$

It is employed to account for endogeneities in the regressors x_{0t} caused by any cointegration between x_{0t} and y_t. The second correction takes into account the effects of serial covariances in the shocks u_{1t} and any serial covariance between u_{ot} and the history of u_{1t}. The bias effect arises from the persistence of shocks due to the unit roots in X_{1t}. The induced one sided long run covariance matrices carry these effects in an OLS regression (Phillips, 1995). They can be defined as

$$\widehat{\Omega}_{0.x} = \widehat{\Omega}_{00} - \widehat{\Omega}_{0x} \widehat{\Omega}_{xx}^{-1} \widehat{\Omega}_{x0}. \tag{28}$$

The correction is then given by

$$\widehat{\Delta}_{ox}^{*} = \widehat{\Delta}_{ox} - \widehat{\Omega}_{0x} \, \widehat{\Omega}_{xx}^{-1} \widehat{\Delta}_{x.x} \tag{29}$$

Combining both corrections the formular for the fully modified estimator is [20]

$$\hat{\beta}^* = (Y^{*\prime} X - T \widehat{\Delta}_{ox}^{*})(X'X)^{-1}. \tag{30}$$

3.3 Regime shifts in Cointegration models

To apply the FM-OLS Estimator in a model with structural changes we make use of the approach developed by Hansen (2003) which allows the parameters to change their values at the break points.[21]

We rewrite equation (22) with $\tau(\mathrm{t})$ as a constant.

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[19] Other Studies adopt the estimator of Newey and West (1987) which is robust to serial correlation and heteroskedasticity. For Details see Cappuccio and Lubian (2003).

[20] The traditional OLs estimator is given by $\hat{\beta} = Y'X(X'X)^{-1}$

[21] We verified with our results with a related approach introduced by Gregory and Hansen (1996). They model the changes in the intercept and the slope coefficients compared to the first subperiod running from 0 to T_1. The base model is than written as $y_t = \tau_1 + \tau(\mathrm{t}) + x_t'\kappa_1 + x_t'\kappa(t) + z_t'\delta_1 + z_t'\delta_j(t) + u_t$.

$$y_t = \tau(\mathrm{t}) + x_t'\kappa(t) + z_t'\delta_j(t) + u_t \quad (31)$$

The piecewise constant time-varying parameters are given by

$$\delta_j(t) = \delta_1 1_{1t} + \cdots \delta_m 1_{mt} \quad (32)$$

$$\kappa_j(t) = \kappa_1 1_{1t} + \cdots \kappa_m 1_{mt} \quad (33)$$

$$\tau(t) = \tau_1 1_{1t} + \cdots \tau_m 1_{mt} \quad (34)$$

where the indicator function for each subsample is defined as follows (Hansen, 2003)

$1_{mt} = 1(T_{j-1} + 1 < t < T_j)$, J=1,…….m

with the convention that T_0=0 and $T_m = T$. Defining dummies according to the indicator function assures that we are able to obtain separate estimates for each period.

4. Data and estimated Models

4.1 Data

Our sample contains monthly data running from January 1975 until December 2007. We use the aggregate M1 for money supply. Real income is proxied by the real production index. As suggested by Wolff (1987) the Producer Price index serves as a proxy for tradeable goods while the basket of non-tradeables is reflected by the CPI. Furthermore, we use the overall trade balance as an approximation of the cumulated current account. Since unit root tests suggest that the cumulated overall trade balance is integrated of order 2 we decided to apply first differences for the US and the EMU series. This can be done without changing the underlying economic theory. As seen in the HP model, the equilibrium flow determines the equilibrium stock. Since the bilateral trade balance can be expressed in two currencies, it is not quite clear which denomination currency should be used. In case of our analysis a separate cointegration analysis (not reported) has shown that dollar denominated balance adjusts to the Euro denominated one. Thus, we decided to choose the Euro configuration. Exchange rates, money supply and real income are expressed in logs. All series have been seasonally adjusted and are taken from International Financial Statistics of the International Monetary Fund.

In sharp contrast with other studies investigating the Euro exchange rate, we rely on the Deutsche Mark and the fundamentals of Germany before the introduction of the Euro. The reason is that we are interested in market rates which would be contrasted by using the ECU. Although the Deutsche Mark had a similar importance on the foreign exchange market measured by its market turnover, we do not see the Deutsche Mark as a predecessor for the Euro. Nevertheless, we use one time series which contains the German values until December 1998 and the values of the EMU further on. Consequently, the Deutsche Mark/ US Dollar exchange rate is converted by the official Deutsche Mark/ Euro exchange rate in order to obtain a level adjustment. As a consequence, we also adjust the German fundamentals in levels to allow for a smooth transition to the EMU data. Since we are dealing with structural break models in the empirical section, we do not see any problems with our proceeding. In the following the term EMU refers to time series which contains German data and the Euro to Deutsche Mark until the beginning of 1999.

4.2 Preliminary tests for unit roots and stationarity

Although the FM-OLS estimator and the Bai Perron methodology are able to handle a combination of I(0) and I(1) regressors, testing the Data for unit roots is necessary as a first step. With the exchange rate being an I(1) variable, the concept of cointegration only makes sense if the fundamentals can also be treated as I(1) processes. By definition, a cointegrating relationship can only exist between variables which are integrated of the same order (Engle and Granger 1987). Neither can a stationary variable force a nonstationary variable to adjust, nor is a stationary relationship between I(1) and I(2) variables possible. Furthermore, reference in a model with I(2) variables is far more complicated from a statistical point of view.

To test for unit roots, we apply the Phillips-Perron Test, the KPSS Test and the DF-GLS Test. In the first stance we test for stationarity in the levels. Afterwards we take differences and check if a unit root remains i.e. if the corresponding variables are integrated of order two. If both hypotheses are rejected we conclude that the variable is I(1). The results are presented in Table 1. According to the results, all variables can be considered as being integrated of order one although the results are mixed in some cases. The KPSS test rejects the hypothesis of stationarity for the change in the money supply of the United States, the change in the bilateral trade balance and the twice differenced trade balance of the euro area. Furthermore, the hypothesis of a unit root is accepted for the change in the trade balance of the euro area according to the DF-GLS test. However, due to the fact that the other tests showed contrary results for these series we treat them as I(1).

4.3 Empirical results

4.3.1 Estimation of the break points

The break points we achieved by applying the Bai and Perron methodology are presented in Table 2. Obviously, breaks occur frequently so we conclude that there is no stable long-run equilibrium relationship among fundamentals and exchange rates since the breakdown of Bretton Woods. Another result is that despite differences some significant similarities between the various configurations remain. The number of break points always lies among eight and ten although we allowed for a shift every twelve months. Furthermore, the location of the dates for the different models is close together. An encouraging result is that major economic or political developments deliver good explanations for instabilities in many cases. The breaks in 1977 and 1978 can clearly be addressed to macroeconomic turbulences arising from the oil price shocks and worldwide recessions. Furthermore, instabilities often occur within the epoch of the so-called pseudo monetarism policy by the FED within 1979 and 1982 or at the end of the rise of the dollar during the mid 1980s. The next date of October 1988 is more difficult to interpret. The election of George Bush and the G 7 summit in Berlin[22] offer possible explanations. For each model, breaks are detected in February of 1992, shortly after the German Reunification. The following instability in 1992 and 1993 might be attributed to the crisis of the European Monetary System. Within a comparatively stable period until the end of the 1990s the only instability in 1997 might have been caused by the Asian Currency Crisis or the deep worsening in the US trade balance which started in 1996. Afterwards, breaks are reported for three models in 2000 and for each model in November of 2004. In the mid of 2000, the American economy started to slow down with the American stock market crashing. Interestingly, the last break exactly coincides with event to which the short-term interest rates of the Euro area declined the below the level of US interest rates. Of course, concerning all these dates we can only guess and many other important developments are not reflected by breaks.

4.3.2 Interpretation of the time varying coefficients

Going one step further we proceed by estimating the cointegration vector via FM-OLS using the obtained break dates. Table 3 contains the results for the specified models. Most of the breaks have occurred at the same time. Since the configuration 1 is embedded in the other three configurations, we predominantly draw on the results of configurations 2 (Table 4), 3 (Table 5) and 4 (Table 6) and use the configuration 1 (Table 3) for comparisons. Our proceeding in the analysis is as follows: we treat model 3 as the main model because this is the model which is mainly used in the empirical literature (e.g. Meese and Rogoff, 1983). For comparisons, we draw on model 2 for the distinction between the overall net foreign (nfa) asset positions of each country (in our case the changes in the nfas) and the bilateral net foreign asset position. A comparison of model 3 and 4 helps us in separating real effects.

Models, 1, 2 and 3 are broadly consistent with the real interest rate model in the first sub-period after our period of observation starts. Only in the case of model 4 the inflation expectations of the EMU enter with an incorrect sign. While the overall change in nfa of the EMU in model 4 is not significant, the variable of the US show significance at the 1% level. The negative sign indicates that risk considerations seem to be important. A worsening in the US trade account is linked to a depreciation of the US Dollar. Important to notice is that the money supply of the USA seem to be strongly linked to the exchange rate. During this period they seem to share common trends with the Euro-US Dollar exchange rate.

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[22] In contrast to previous meetings, the participants of Berlin did not publically claim that Fluctuations in the Dollar are unwanted

From 1977:05 till 1979:12 many coefficients of model 3 show signs which are not consistent with standard theory. Both the EMU money supply and the EMU inflation expectations are highly significant with a negative sign. When either the relative price of tradables or the bilateral nfa is taken into account, they show the correct sign and the significance of the money supply and the inflation rates in model 2 and 4 disappear. On the one hand the reason is that the sub-periods of model 2 and 4 are similar in this example but different from model 3. On the other hand the second oil price shock lies in this period. It becomes obvious that real factors have an impact on the exchange rate and let the impact of nominal factors vanish. The period from 1979:12 till 1981:06 in model 3 is again broadly consistent with the theory. The only deviation from the RID model is that the EMU short-term interest rates enter with a positive sign which indicate that either opportunity cost of holding money are important in the short-run or the monetary policy dominates such that the high correlation between exchange rate and short-term interest rates arises. Between 1981 and the end of 1984 the US money supply and the US real income show signs which are not consistent with standard theory. However, the inclusion of the change of the overall US nfa and the relative prices of tradables yield signs which are consistent with the theory. Only the bilateral nfa have a positive sign which means that an increase of EMU claims on US assets coincide with a depreciation of the Euro. Such a correlation can be explained by anticipated monetary shocks and occur during the overshooting period. In this period the US Dollar appreciated strongly against major currencies. During the overshooting period after an announced monetary expansion the currency appreciates while the money stock widens. The money inflow generates current account deficits. This linkage is reflected in the positive sign of the bilateral nfa.

The following period (1984:07-1988:10 for model 1 and – 1988:08 for model 3, 1985:03-1988:10 for model 2 and 4) is characterized by interventions which should weaken the US Dollar.[23] In all models, inflation expectations in the EMU are highly significant while US real income show mostly an incorrect sign based upon standard theory which is predominantly due to the interventions occurred. In the next period, which starts in 1988:10 (except for model 3 in 1988:08) and ends in 1991:02, all signs are broadly consistent with the theory. The results indicate that liquidity effects are important. After the reunification of Germany which seems to be responsible for the next regime, the results of model 2 and 3 give evidence that capital flows and inflation rate expectations are important. Only in model 4 EMU money supply and real income enter significantly whereas the money supply has the incorrect sign. The consideration of asset positions as in model 2 and 3 seems to absorb this effect because it is not significant in these models. After the crises of the European Monetary

[23] The Plaza agreement should depress the US Dollar while in the Louvre accord the depreciating tendency of the US Dollar should be stopped.

System, the next sub-periods start in 1993:10 (model 2) or 1993:12 (model 3 and 4) and end differently. In model 2 the next regime starts in 2000:01, in model 3 in 1997:06 and in model 4 in 1999:03. As a consequence, the results of the different models vary remarkably. The only analogy can be observed with respect to inflation rate expectations. They seem to be of equal importance. On the one hand this result is not surprising because the durations of the regimes are not equal. On the other hand, these are the longest sub-periods for model 2 and 4 and we would have expected that the coefficients and their signs are similar.

However, the inclusion of either bilateral nfa, overall nfa or relative prices of tradables changes the results considerably. For model 1, 2 and 3 a further regime starts during 2000 (in 2000:01 for model 1 and 3 and 2000:07 for model 2) and for model 4 at the beginning of 1999. In addition, model 3 generates an additional break in 1997:06. The period between the end of 1993 and the beginning of 2000 is absolutely not compatible with standard theory. A reason for this additional break can be seen in the use of the changes in overall nfa. Since the changes of overall nfa are simply equal to the current account balance. It is widely known that the US current account started to widen in mid 1997. This might be the reason why we obtain these results from our analysis. Consequently, the change in the US current account dominated the effects.

The first years of the Euro also yield results for both the EMU and the US money supply and real income which show signs contrary to standard theory. Nevertheless, the relative prices of tradables have the correct signs. From this point of view, real effects had an important impact during this period. In the last regime which is equal to within four models this seems to be characterized by overshooting in the sense of Frankel (1979).

A clear impact of net foreign asset positions cannot be stated. Both the accumulation of overall net foreign assets and the bilateral net foreign asset position are not significant in every regime. In the periods in which they are significant the sign changes frequently. Nevertheless, there is only one period in which the change in overall nfa has the same signs. This is from 1997:06 till 2000:01.

In model 4 all coefficient of the US foreign prices have the same sign. i.e. an increase in US relative price of tradables results in a depreciation of the US Dollar. For the Euro series only during the period from March 1999 to November 2004, after the introduction of the Euro, the coefficient has not the correct sign. Taken together the nominal exchange rate is linked to US relative prices in five periods. From this point of view, it can be said based upon the results of model 4 that the nominal exchange rate is only correlated with real variables in five periods which show a concentration in two periods of time. These two periods run from beginning of 1976 to the beginning of 1985 (before the interventions started) and from the beginning of 1991 to the end of 2004. The remaining periods (1975:01-1976:12, 1985:03-1991:02) are characterized by financial distress and interventions. During 2004:1 and 2007:12 inflation expectations concerning the USA became more important and as a consequence the relative price of tradables is less important.

Finally, we can conclude that the relationship between exchange rate and fundamentals over a period of at least one and a half year is stable (otherwise the Bai-Perron Test would have estimated more breaks as our configuration allows for breaks every 12 months). However, the linkage between exchange rate and fundamentals differ in each period.

4.3.3 Analysis of the error correction term

In the last part of our analysis we examine if the implicit assumption of cointegration is valid. According to the various unit root tests reported in Table 8 the error term resulting from the step wise relationships should be considered as stationary which gives clear evidence in favour of long-run relationship between exchange rates and fundamentals. Furthermore, a regression of the exchange rate's change on the error term shows that the exchange rate is significant and enters with a negative coefficient. The corresponding results which are summarized in the Tables 9-12 show that this true except for the first period of model 2 which only lasts 8 months. This is an indication for an error correcting behaviour, meaning the exchange rate adjusts to disequilibria. An interesting question is if this correction mechanism is also a subject to structural changes. To tackle this question we apply the Bai and Perron test once again but in the following without a restriction on the minimum distance between two breaks. The results which are summarized in Table 7 show that we observe four break points for model 2 and three nearly equal break points for the other models. Hence, we conclude that structural breaks in the cointegration coefficients are more frequent than in the adjustment coefficients. Again, the location of the breaks can be associated with economic developments. The first break point in model 3 can again be addressed to the rising oil price. The explanations for the breaks in the cointegrating coefficients can also be applied to 1980 and 1985. Surprisingly, the last break point occurs in 1987 with the Louvre accord as a possible cause.

5. Conclusion

In this paper, we have examined the long run relationship between exchange rates and fundamentals with respect to structural breaks in the coefficients. For the Euro-US Dollar we can show that fundamentals are important in each subperiod we obtained by our analysis but their impact differs significantly among the different regimes. With respect to this issue we can draw some major conclusions. Firstly, there is no regime in which no fundamentals enter. Furthermore, there are no perseverative regimes, i.e. either the coefficient values for the same fundamentals differ or the significance differs. Our results contradict the view that fundamentals only matter during single periods while having no explanatory content within other regimes. We also verified the assumption that a cointegrating relation exists by testing the error terms for stationarity. Moreover, the exchange rate adjusts to the step-wise linear relationships in all cases. Altogether a linear modeling of exchange rates is inadequate in many cases. Another result is that economic developments can be consulted to explain the date of the breaks as well as the sign and the significance of the parameters in many cases. This topic surely needs further attention. We also leave the exercise to verify our results with regard to other currencies or model configurations for further research.

Appendix

Table 1: Unit Root tests

	Levels				First Differences			
	PP	DF-GLS		KPSS	PP	DF-GLS		KPSS
	test statistic[a]	lags	test statistic[b]	test statistic[c]	test statistic[a]	lags	test statistic[b]	test statistic[c]
USD/EURO	-1.317	2	-0.437	2.690**	-16.66**	0	-1.485	0.084
i_s^{emu}	-1.970	0	-1.154	1.840**	-19.86**	0	-17.069*	0.0743
P^{emu}	-2.594	12	-0.651	2.012**	-17.32**	0	-7.782**	0.110
ΔCTB^{EMU}	-4.048*	0	-4.643*	0.566*	-31.772*	0	-30.161*	0.062
i_s^{us}	-1.899	12	-1.636	3.466*	-16.559*	0	-16.480*	0.456
P^{emu}	-2.581	12	-0.373	3.551*	-13.701*	0	-13.606*	0.178
M^{emu}	-1.662	0	-1.691	1.008*	-21.800*	0	-19.335*	0.123
Y^{emu}	-3.360	15	-2.693	0.182**	-31.059*	0	-25.513*	0.049
M^{us}	-0.027	8	-0.669	1.543*	-15.202*	16	-2.121**	1.696*
Y^{us}	-1.839	0	-1.253	0.489*	-15.268*	0	-3.335*	0.083
ΔCTB^{us}	-0.620	0	-0.974	1.336*	-28.596*	0	-16.376*	0.628**
BCTB	-2.383	15	-1.754	0.419*	-4.446*	0	-3.012*	1.210*

Note: * Statistical significance at the 5% level, ** at the 1% level. For the PP test and the DF-GLS test the series contain a unit root under the null whereas the KPSS test assumes stationarity under the null. [a] Critical values are taken from MacKinnon (1991): 5% -2.86, 1% -3.43. [b] Critical values are given by Elliot et al. (1996): 5% -1.95, 1% -2.58. Number of lag is chosen by using the modified AIC (MAIC) by Ng/Perron (2001). Maximum lag number is chosen by Schwert (1989) criterion. [c] Critical values are given by Kwiatkowski et al. (1992): 5% 0.463, 1% 0.739. Autocovariances are weighted by Bartlett kernel.

Table 2: Comparison of Breaks

	Model 1	Model 2	Model 3	Model 4
	1975:01	1975:01	1975:01	1975:01
	1977:01	1977:04	1977:05	1976:12
	1980:02		1979:12	
	1981:06	1981:06	1981:06	1981:09
	1984:07		1984:07	
		1985:03		1985:03
	1988:10	1988:10	1988:08	1988:10
	1991:02	1991:02	1991:02	1991:02
	1992:10	1993:10	1993:12	1993:12
	1995:02			
			1997:06	
				1999:03
	2000:01	2000:07	2000:01	
	2004:11	2004:11	2004:11	2004:11
	2007:12	2007:12	2007:12	2007:12
No. of breaks	10	8	10	8

Note: Breaks within a horizon of 6 month a seen as comparable.

Table 3: Estimation results of model 1

	c	m^{EMU}	y^{EMU}	i_s^{EMU}	p^{EMU}	m^{US}	y^{US}	i_s^{US}	p^{US}
1975:01	12.010 ***	-0.188	-0.494 *	-0.042 ***	2.551	-1.997 ***	0.803	0.030 ***	-3.581 ***
	0.000	*0.353*	*0.077*	*0.000*	*0.108*	*0.000*	*0.153*	*0.000*	*0.000*
1977:01	-9.789 ***	-0.898 ***	0.098	0.017 ***	0.342	-0.229	1.072 ***	-0.013 ***	-2.005 ***
	0.001	*0.003*	*0.613*	*0.004*	*0.731*	*0.630*	*0.002*	*0.000*	*0.008*
1980:02	-2.721	0.647 *	-0.272	0.342	3.239 *	-2.497 ***	0.883	0.006	-8.454 ***
	0.751	*0.075*	*0.635*	*0.731*	*0.061*	*0.006*	*0.163*	*0.206*	*0.000*
1981:06	-12.664 ***	-0.365	-0.421 ***	0.047 ***	-5.920 ***	1.023 ***	-0.349	-0.004 *	-0.321
	0.000	*0.203*	*0.001*	*0.000*	*0.000*	*0.003*	*0.216*	*0.097*	*0.652*
1984:07	-0.570	-0.213	0.014	0.077 ***	8.270 ***	-0.320 **	-2.020 ***	-0.004	-0.701
	0.840	*0.321*	*0.940*	*0.000*	*0.000*	*0.016*	*0.000*	*0.338*	*0.227*
1988:10	18.732 ***	0.301 ***	-0.079	0.015 **	0.639	-5.006 ***	0.198	0.011	-2.868 ***
	0.000	*0.001*	*0.786*	*0.011*	*0.681*	*0.000*	*0.609*	*0.254*	*0.001*
1991:02	-9.267	-2.980 ***	0.132	-0.051 **	4.270 ***	1.348 **	1.902 **	0.053 ***	-0.119
	0.134	*0.001*	*0.773*	*0.012*	*0.000*	*0.047*	*0.026*	*0.010*	*0.912*
1992:10	-14.034 ***	-0.552	-0.089	0.004	-0.295	0.562	0.612	-0.042 ***	-8.629 ***
	0.005	*0.339*	*0.780*	*0.841*	*0.777*	*0.313*	*0.321*	*0.004*	*0.000*
1995:02	-16.266 ***	0.282	0.577 **	-0.071 ***	6.031 ***	-0.185	0.145	0.067 ***	-6.935 ***
	0.000	*0.130*	*0.017*	*0.000*	*0.000*	*0.221*	*0.627*	*0.000*	*0.000*
2000:01	-2.319	-0.870 ***	0.154	0.069 ***	7.887 ***	0.725 ***	-1.996 ***	0.001	-2.134 ***
	0.476	*0.000*	*0.665*	*0.000*	*0.000*	*0.005*	*0.000*	*0.922*	*0.000*
2004:11	-17.172 **	0.727 ***	-0.520	-0.103 ***	0.039	0.244	-0.016	0.019 *	-1.339
2007:12	*0.012*	*0.000*	*0.328*	*0.000*	*0.986*	*0.733*	*0.979*	*0.052*	*0.155*

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 4: Estimation results of model 2

	c	m^{EMU}	y^{EMU}	i_s^{EMU}	p^{EMU}	m^{US}	y^{US}	i_s^{US}	p^{US}	$BCTB$
1975:01	3.588	0.060	-0.402	-0.041 ***	-1.640	-0.814	1.002 *	0.027 ***	-1.896 ***	0.031 *
	0.191	*0.765*	*0.172*	*0.000*	*0.292*	*0.238*	*0.069*	*0.000*	*0.001*	*0.077*
1977:04	-7.850 **	0.050	0.177	0.012 ***	0.523	-0.649 ***	1.260 ***	-0.004 **	-1.968 ***	-0.066 ***
	0.011	*0.832*	*0.327*	*0.000*	*0.508*	*0.008*	*0.000*	*0.046*	*0.000*	*0.000*
1981:06	-3.582	0.450	-0.255 **	0.058 ***	-4.730 ***	0.517	-0.953 ***	-0.003	-1.275 *	0.028 ***
	0.245	*0.144*	*0.039*	*0.000*	*0.000*	*0.157*	*0.000*	*0.349*	*0.065*	*0.000*
1985:03	-10.473 *	0.522	0.316	0.068 ***	7.894 ***	0.462	-0.337	-0.014 *	-0.798	-0.028 ***
	0.074	*0.108*	*0.188*	*0.000*	*0.000*	*0.103*	*0.631*	*0.084*	*0.255*	*0.000*
1988:1	34.564 ***	0.230 **	-0.504	-0.009	1.709	-5.288 ***	-0.539	0.017	-3.644 ***	0.032 *
	0.000	*0.020*	*0.177*	*0.504*	*0.313*	*0.000*	*0.323*	*0.126*	*0.000*	*0.055*
1991:02	-2.419	-0.460	-0.175	-0.044 ***	3.579 ***	0.624	-0.317	0.071 ***	-1.173	-0.058 ***
	0.532	*0.383*	*0.435*	*0.004*	*0.000*	*0.173*	*0.664*	*0.000*	*0.229*	*0.007*
1993:1	-17.710 ***	1.467 ***	0.589 ***	0.040 ***	2.572 ***	-0.861 ***	1.735 ***	-0.032 ***	-4.682 ***	-0.065 ***
	0.000	*0.000*	*0.007*	*0.000*	*0.001*	*0.000*	*0.000*	*0.000*	*0.000*	*0.000*
2000:07	26.398 ***	-2.711 ***	-0.252	0.073 ***	6.702 ***	0.170	-2.359 ***	-0.022 **	-0.687	0.053 ***
	0.000	*0.000*	*0.544*	*0.000*	*0.000*	*0.578*	*0.000*	*0.021*	*0.304*	*0.000*
2004:11	-5.235	1.362 ***	-0.661	-0.014	1.290	-0.745	0.012	0.025 **	-1.535	-0.030 ***
2007:12	*0.476*	*0.000*	*0.244*	*0.692*	*0.592*	*0.370*	*0.985*	*0.017*	*0.127*	*0.005*

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 5: Estimation results of model 3

	c	m^{EMU}	y^{EMU}	i_s^{EMU}	p^{EMU}	m^{US}	y^{US}	i_s^{US}	p^{US}	ΔCTB^{EMU}	ΔCTB^{US}
1975:01	9.109***	-0.088	-0.447**	-0.038***	-0.475	-1.729***	0.913**	0.032***	-2.883***	0.008	-0.014***
	0.000	*0.552*	*0.031*	*0.000*	*0.678*	*0.000*	*0.016*	*0.000*	*0.000*	*0.407*	*0.004*
1977:05	-3.987**	-1.536***	-0.042	0.025***	-3.207**	0.536	0.325	-0.014***	-0.635	0.000	0.004
	0.049	*0.000*	*0.810*	*0.000*	*0.013*	*0.249*	*0.348*	*0.000*	*0.393*	*0.992*	*0.277*
1979:12	3.808	1.391***	-1.008***	0.016***	7.810***	-3.115***	0.502	0.008**	-8.485***	0.026***	0.019***
	0.567	*0.000*	*0.009*	*0.000*	*0.000*	*0.000*	*0.217*	*0.014*	*0.000*	*0.007*	*0.000*
1981:06	-8.491***	-0.169	-0.441***	0.047***	-5.689***	0.645**	-0.375*	-0.002	-0.893	0.004	-0.006***
	0.000	*0.460*	*0.000*	*0.000*	*0.000*	*0.018*	*0.085*	*0.291*	*0.114*	*0.458*	*0.000*
1984:07	0.256	-0.098	0.309**	0.081***	7.791***	-0.317***	-2.002***	-0.001	-1.140**	-0.018***	-0.001
	0.876	*0.566*	*0.046*	*0.000*	*0.000*	*0.006*	*0.000*	*0.773*	*0.014*	*0.000*	*0.309*
1988:08	22.217***	0.239***	-0.291	0.011**	0.351	-5.025***	0.453	0.001	-2.383***	-0.005	0.004*
	0.000	*0.001*	*0.177*	*0.023*	*0.769*	*0.000*	*0.214*	*0.874*	*0.001*	*0.188*	*0.084*
1991:02	-10.491***	-0.291	-0.162	-0.024***	4.754***	0.207	0.577	0.046***	-0.298	-0.013***	0.006**
	0.000	*0.403*	*0.229*	*0.000*	*0.000*	*0.487*	*0.208*	*0.000*	*0.661*	*0.000*	*0.010*
1993:12	-25.356***	1.142***	-0.034	0.072***	6.895***	0.601	0.852*	-0.051***	-6.874***	0.004	0.001
	0.000	*0.000*	*0.881*	*0.000*	*0.000*	*0.198*	*0.063*	*0.000*	*0.000*	*0.148*	*0.593*
1997:06	-16.296***	-1.003***	0.902***	-0.058***	0.520	2.450***	-1.541***	0.053**	-2.036	-0.007***	-0.014***
	0.000	*0.000*	*0.003*	*0.000*	*0.657*	*0.000*	*0.001*	*0.022*	*0.282*	*0.002*	*0.000*
2000:01	6.302***	-0.923***	0.122	0.059***	7.657***	0.446**	-2.729***	0.013**	-2.205***	0.005***	-0.004***
	0.004	*0.000*	*0.664*	*0.000*	*0.000*	*0.028*	*0.000*	*0.036*	*0.000*	*0.000*	*0.000*
2004:11	-13.322**	0.733***	-0.462	-0.099***	-0.057	0.189	-0.200	0.016*	-1.493**	-0.001	***
2007:12	*0.011*	*0.000*	*0.260*	*0.000*	*0.974*	*0.732*	*0.700*	*0.059*	*0.043*	*0.455*	

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 6: Estimation results of model 4

	c	m^{EMU}	y^{EMU}	i_s^{EMU}	p^{EMU}	m^{US}	y^{US}	i_s^{US}	p^{US}	$\left(\frac{P^T}{P^{NT}}\right)^{EMU}$	$\left(\frac{P^T}{P^{NT}}\right)^{US}$
1975:01	16.103***	-0.257	-0.631*	-0.042***	-5.103**	-2.362***	0.604	0.030***	-3.920***	-0.020	0.006
	0.000	*0.232*	*0.056*	*0.000*	*0.018*	*0.000*	*0.305*	*0.000*	*0.000*	*0.283*	*0.852*
1976:12	-15.259***	0.104	-0.188	0.011***	0.482	-0.982***	1.187***	-0.001	-4.012***	0.085***	-0.015***
	0.000	*0.637*	*0.224*	*0.001*	*0.470*	*0.001*	*0.000*	*0.546*	*0.000*	*0.000*	*0.009*
1981:09	-20.160***	-0.313	0.184	0.027***	0.007	0.794**	0.195	-0.001	-2.423***	0.034***	-0.044***
	0.000	*0.488*	*0.129*	*0.002*	*0.995*	*0.011*	*0.526*	*0.797*	*0.001*	*0.002*	*0.000*
1985:03	-4.492	-0.102	0.253	0.069***	8.228***	-0.497**	-2.185***	-0.003	-1.342*	-0.012	0.009
	0.262	*0.673*	*0.270*	*0.000*	*0.000*	*0.030*	*0.000*	*0.675*	*0.056*	*0.120*	*0.248*
1988:1	15.427***	0.250***	-0.152	0.013**	2.395	-5.127***	0.394	0.018*	-1.696	-0.012	-0.011
	0.001	*0.006*	*0.602*	*0.046*	*0.159*	*0.000*	*0.412*	*0.077*	*0.162*	*0.140*	*0.158*
1991:02	-13.683***	-0.917*	-0.419**	-0.021*	3.828***	0.473	0.546	0.065***	0.882	0.009	-0.031**
	0.002	*0.056*	*0.034*	*0.057*	*0.000*	*0.280*	*0.372*	*0.000*	*0.379*	*0.438*	*0.011*
1993:12	3.215	-0.299	-0.095	0.001	7.240***	-2.238***	-0.256	-0.003	-6.018***	-0.016***	-0.018***
	0.503	*0.259*	*0.673*	*0.955*	*0.000*	*0.000*	*0.288*	*0.714*	*0.000*	*0.003*	*0.000*
1999:03	-12.012***	-0.375**	1.001***	0.053***	7.198***	0.434*	-2.019***	0.017**	-0.443	0.015***	-0.017***
	0.003	*0.034*	*0.005*	*0.000*	*0.000*	*0.062*	*0.000*	*0.015*	*0.520*	*0.003*	*0.000*
2004:11	-20.872***	0.529***	-0.366	-0.107***	1.376	0.201	0.143	0.032**	-2.335**	-0.005	0.005
2007:12	*0.004*	*0.004*	*0.493*	*0.000*	*0.551*	*0.783*	*0.815*	*0.024*	*0.032*	*0.424*	*0.107*

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 7: Comparison of Breaks in the error correction model

	Model 1	Model 2	Model 3	Model 4
	1975:01	1975:01	1975:01	1975:01
		1975:09		
	1980:07	1980:02	1980:01	1980:07
	1985:03	1985:03	1985:03	1985:03
	1987:02	1987:02	1987:02	1987:02
	2007:12	2007:12	2007:12	2007:12
No. of breaks	3	4	3	3

Note: Breaks within a horizon of 6 month are seen as comparable.

Table 8: Unit root tests for the error terms

	Levels			
	PP	DF-GLS		KPSS
	test statistic[a]	lags	test statistic[b]	test statistic[c]
Model 1	-16.20***	2	-14.60***	0.013
Model 2	-14.67***	0	-14.54***	0.013
Model 3	-18.50***	0	17.54**	0.013
Model 4	-15.06***	0	-14.34***	0.027

Note: * Statistical significance at the 5% level, ** at the 1% level. For the PP test and the DF-GLStest the series contain a unit root under the null whereas the KPSS test assumes stationarity under the null.[a] Critical values are taken from MacKinnon (1991): 5% -2.86, 1% -3.43. [b] Critical values are given by Elliot et al. (1996): 5% -1.95, 1% -2.58. Number of lag is chosen by using the modified AIC (MAIC) by Ng/Perron (2001). Maximum lag number is chosen by Schwert (1989) criterion. [c] Critical values are given by Kwiatkowski et al. (1992): 5% 0.463, 1% 0.739. Autocovariances are weighted by Bartlett kernel.

Table 9: Error correction estimation of model 1

	Constant	Coefficient
1975:01	-0,0044	-0,3401**
	0,1041	*0,0264*
1980:07	0,0157***	-0,5583***
	0,003	*0,0005*
1985:03	-0,0226***	-0,6647***
	0,0000	*0,0000*
1987:02	0,0032	-0,3874***
	0,3345	*0,0000*

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 10: Error correction estimation of model 2

	Constant	Coefficient
1975:01	0,0161**	-0,2256
	0,0295	*0,5539*
1975:09	-0,0236	-0,5524***
	0,0018	*0,0000*
1980:07	-0,0054	-0,7063***
	0,4991	*0,0000*
1985:03	-0,042***	-0,6168***
	0,0000	*0,0062*
1987:02	-0,017***	-0,361***
	0,0229	*0,0000*

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 11: Error correction estimation of model 3

	Constant	Coefficient
1975:01	-0,0042	-0,3080***
	0,1237	*0,0027*
1980:07	0,0156***	-0,6771***
	0,0005	*0,0001*
1985:03	-0,022***	-0,6554***
	0,0000	*0,0000*
1987:02	0,0029	-0,3833***
	0,3763	*0,0000*

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 12: Error correction estimation of model 4

	Constant	coefficient
1975:01	-0,0052	-0,2515*
	0,0755	*0,0538*
1980:07	0,0158***	-0,734***
	0,0003	*0,0000*
1985:03	-0,022***	-0,717***
	0,0000	*0,0000*
1987:02	0,0041	-0,452***
	0,2397	*0,0000*

Note: p-values are in italic letters. * Statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

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