

Beckmann, Joscha; Belke, Ansgar; Kühl, Michael

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How Stable Are Monetary Models of the Dollar-Euro Exchange Rate? - A Time-varying Coefficient Approach

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Joscha Beckmann, Ansgar Belke,
and Michael Kühl

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Technische Universität Dortmund, Department of Economic and Social Sciences
Vogelpothsweg 87, 44227 Dortmund, Germany

Universität Duisburg-Essen, Department of Economics
Universitätsstraße 12, 45117 Essen, Germany

Rheinisch-Westfälisches Institut für Wirtschaftsforschung (RWI)
Hohenzollernstr. 1/3, 45128 Essen, Germany

Editors:

Prof. Dr. Thomas K. Bauer
RUB, Department of Economics
Empirical Economics
Phone: +49 (0) 234/3 22 83 41, e-mail: thomas.bauer@rub.de

Prof. Dr. Wolfgang Leininger
Technische Universität Dortmund, Department of Economic and Social Sciences
Economics – Microeconomics
Phone: +49 (0) 231 /7 55-32 97, email: W.Leininger@wiso.uni-dortmund.de

Prof. Dr. Volker Clausen
University of Duisburg-Essen, Department of Economics
International Economics
Phone: +49 (0) 201/1 83-36 55, e-mail: vclausen@vwl.uni-due.de

Prof. Dr. Christoph M. Schmidt
RWI
Phone: +49 (0) 201/81 49-227, e-mail: christoph.schmidt@rwi-essen.de

Editorial Office:

Joachim Schmidt
RWI, Phone: +49 (0) 201/81 49-292, e-mail: joachim.schmidt@rwi-essen.de

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Joscha Beckmann, Ansgar Belke, and Michael Kühl*

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Abstract

This paper examines the significance of different fundamental regimes by applying various monetary models of the exchange rate to one of the politically most important exchange rates, the exchange rate of the US dollar vis-à-vis the euro (the DM). We use monthly data from 1975:01 to 2007:12. Applying a novel time-varying coefficient estimation approach, we come up with interesting properties of our empirical models. First, there is no stable long-run equilibrium relationship among fundamentals and exchange rates since the breakdown of Bretton Woods. Second, there are no recurring regimes, i.e. across different regimes either the coefficient values for the same fundamentals differ or the significance differs. Third, there is no regime in which no fundamentals enter. Fourth, the deviations resulting from the stepwise cointegrating relationship act as a significant error-correction mechanism. In other words, we are able to show that fundamentals play an important role in determining the exchange rate although their impact differs significantly across different sub-periods.

JEL Classification: E44, F31, G12

Keywords: Structural exchange rate models, cointegration, structural breaks, switching regression, time-varying coefficient approach

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* Joscha Beckmann, University of Duisburg-Essen; Ansgar Belke, University of Duisburg-Essen and IZA Bonn; Michael Kühl, University of Goettingen. – We are grateful for valuable comments from the participants of the XII. Applied Economics Meeting Conference, June 4–6, 2009, Madrid, and the participants of the 12th Goettinger Workshop for International Economic Relations, March 12–14, 2009, Goettingen. – All correspondence to Ansgar Belke, University of Duisburg-Essen, Department of Economics, 45177 Essen, Germany, e-mail: ansgar.belke@uni-due.de.

1. Introduction

Disentangling the main drivers of exchange rates 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 (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 past exchange rates with the help of fundamentals and perform poorly in forecasting, in particular (Meese and Rogoff, 1983 and 1988). The results of the seminal study by Meese and Rogoff (1983) still represent the benchmark: exchange rate forecasts by structural models can hardly outperform naïve random walk forecasts (Rogoff, 2009).

Since then many contributions have tried to refute their results. Sticking to the implicit assumption that exchange rates and fundamentals are cointegrated and implementing exogenous parameter restrictions, a couple of authors find predictability in the long run for a similar period as in Meese and Rogoff (Mark, 1995; Chinn and Meese, 1995).¹ However, extending the estimation period yields mostly contrary findings (Kilian, 1999; Abhyankar, Sarno and Valente, 2005). A critical point is the implicit assumption of cointegration which leads to biased conclusions if a stable long run relation does not exist (Berkowitz and Giorgianni, 2001).

While the models of the late 1980s mostly neglect the potential existence of 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 empirical models which are based upon cointegration relationships can indeed improve the evidence in favour of predictability in the long run when periods up to the end of the 1990s are covered (MacDonald and Taylor, 1993, 1994).² However, any extension of the sample period typically yields a breakdown in cointegration relationships (Groen, 1999). Surprisingly, less attention is directed to a closer examination of the link between exchange rates and fundamentals with respect to structural changes in cases where 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 performance. Bacchetta and Wincoop (2009) argue that large and frequent variations in the relationship between the exchange rate and macro fundamentals naturally develop when structural parameters in the economy are unknown and subject to changes. Goldberg and Frydman

¹ Mark (1995) is the first author who focuses on more than one exchange rates simultaneously. He includes the Canadian dollar, the Deutschmark, the Japanese yen and the Swiss franc expressed in US dollar. Chinn and Meese (1995) do include the pound sterling in US dollars as well as the US dollar and the Deutschmark in Japanese yen but not the Swiss franc.

² MacDonald and Taylor (1994) investigate the pound sterling-US dollar exchange rate.

(1996a, 2001) provide evidence that some periods exist in which the monetary model is valid and some other periods in which this is not the case. Thus, the instability of the monetary model in the data generating process might serve as an explanation for the findings of Cheung et al. (2005) which suggest that model specifications that work well in one period do not necessarily work well in another period.³

In the recent past, models capable of taking different regimes into account have been applied to the monetary approach. Sarno, Valente and Wohar (2004) use a Markov regime-switching model in order to investigate the response of exchange rates to deviations from fundamental values in different regimes. Sarno and Valente (2008) demonstrate that exchange rate models that optimally use the information in the fundamentals change often and this implies frequent shifts in the parameters. De Grauwe and Vansteenkiste (2007) investigate particularly the adjustment of the nominal exchange with respect to changes in the fundamentals 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 exchange rate adjustments towards equilibrium paths is nonlinear. To be more specific, fundamentals become important if the deviation from an equilibrium rate is large.

Frömmel, MacDonald and Menkhoff (2005a,b) test directly for the significance of different regimes in the exchange rate determination equation of the real interest rate differential model. The latter is one of the rare but meritorious contributions using a model in which the coefficients in the exchange rate determination process itself are allowed to change. However, since the authors specify their model in first differences, they do not investigate a long-run relationship in a strict sense.⁴ All other contributions focus on deviations of the exchange rate 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 perseverative, i.e. regularly recurring, regimes. In early works, Schinasi and Swamy (1989) and Wolff (1987) apply a time-varying coefficient model (TVP) to monetary models. They are able to show that their models display quite better forecasting properties than fixed coefficient models. Hence, taking into account time-varying parameters appears to be a worthwhile next step towards a valid empirical model of the exchange rate.

³ See also Bacchetta and Wincoop (2009). Parameter instability, i.e. an unstable relationship between exchange rates and macro fundamentals, is confirmed by formal econometric evidence delivered by Rossi (2005).

⁴ In order to obtain a long-run perspective, Frömmel, MacDonald and Menkhoff use annual changes starting from a monthly data set.

Different market surveys suggest that different fundamentals are important during different periods (Gehrig and Menkhoff, 2006, and Christopoulos and Léon-Ledesma, 2009). This pattern can also be derived from the imperfect knowledge approach which is based on the awareness that market participants intermittently revise their views on how fundamentals influence the exchange rate (Frydman and Goldberg, 1996b, 2007). Hence, it is reasonable to assume that a strong and significant relationship between exchange rates and fundamentals exists during sub-periods and that its nature tends to change considerably over time. From this point of view, a fundamental value of the exchange rate exists 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.

Taking these considerations as a starting point, we test for the significance of a couple of different hypotheses in this dynamic context. First, we check whether there has been a stable long-run equilibrium relationship among fundamentals and the US dollar exchange rate vis-à-vis the DM/euro since the breakdown of Bretton Woods I. Second, we test whether the regimes are not perseverative (DEF) implying that across different regimes either the coefficient values for the same fundamentals or the significance differ(s). Third, we check empirically whether there is at least one regime into which no fundamentals enter. Fourth, we test whether the deviation from the stepwise relationship acts as an error-correction term.

The remainder of this paper is organized as follows: Section two introduces the concept of regime-sensitive cointegration and gives a short overview of the models we consider later on. The econometric methodology is described in Section three. We start with a multiple structural change model developed by Bai and Perron (1998, 2003) which we apply to the reduced form of structural exchange rate models. Hypothesis 1 can be rejected if at least one structural change is found. As a next step, we use the estimated breakpoints to generate indicator functions based on which 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 deal with non-stationary variables as regressors and regressands. The results are then evaluated with respect to the second and third hypothesis in Section four. Finally, we construct an error-correction term and regress the change of the exchange rate on this error-correction term in order to investigate whether the exchange rate adjusts to deviations from a fundamental equilibrium relationship.

2. Monetary models of the exchange rate

2.1 Theories

After the breakdown of Bretton Woods I, exchange rate models were developed which see exchange rates as asset prices (Dornbusch, 1976a; Frenkel, 1976; Kouri, 1976). All models of this kind have in common that they rely on a stable money demand function of the form

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

with M the money supply, p the price level and L the money demand depending 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 difference in price levels which is equal to the difference between domestic and foreign money supply less real money demand based on money market equations, so that the exchange rate is determined as follows:

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

In the literature, this model is widely known as the Frenkel and Bilson (FB) model.⁵ In the original monetary model α is zero and $\beta_1 = \beta_1^f = 1$ due to the structure of the money demand function. Equation (2) can be rewritten under the restriction that the (semi-) elasticities of the interest rates are equal. This yields:

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

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 the PPP, the differences in interest rates can then be replaced by the differences in expected rates of inflation.⁶ 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 $E_t(s_{t+1} - s_t) = -\phi(s_t - \bar{s}) + \pi_t - \pi_t^f$.⁷ The real interest rate model (RID) by Frankel (1979) arises if the expectation formation process is combined with the UIP and is solved for the expected change in the exchange rate (equation (4)).

⁵ $\beta_{1,2,3}$ 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.

⁶ This formulation is equivalent to a money demand function in which the expected rates of inflation enter as opportunity costs.

⁷ ϕ denotes the adjustment speed towards the equilibrium value \bar{s} . π is the expected rate of inflation.

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

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 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) to the word, β_1 and β_1^f must be equal to one.⁸ 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 is assumed to be constant in the long-run. 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 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 to accumulate (or to decumulate) 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 many 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 with the cumulated current account. Thus, equation (4) can be extended by the cumulated trade balances as a proxy for the current account balance (eq. (5)).⁹

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

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

⁸ Nevertheless, Driskill and Sheffrin (1981) show that overshooting requires $\beta_1 > 0$ and $\beta_1^f < 0$.

⁹ 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.

insufficient holding of international reserve and foreign indebtedness. A fall in the net foreign asset position (in particular if it is negative) increases the risk premium from which an increase in the exchange rate follows. Hence, the risk premium sensitively reacts 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)).

$$s = \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f - \beta_3 (i_t - i_t^f) + \beta_4 (\pi_t - \pi_t^f) - \beta_5 BCTB_t.^{10} \quad (6)$$

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 (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

$$s = \alpha + \beta_1 m - \beta_1^f m^f - \beta_2 y + \beta_2^f y^f - \beta_3 (i_t - i_t^f) + \beta_4 (\pi_t - \pi_t^f) + \beta_6 \frac{P_t^T}{P_t^{NT}} - \beta_6^f \frac{P_t^{Tf}}{P_t^{NTf}}.^{11} \quad (7)$$

The proportion of traded to non-traded goods mirrors the real exchange rate. A rise in the price of tradeables relative to the price of non-tradeables lets the nominal exchange rate 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 to 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. Hence, the restrictions $\beta_1 = \beta_1^f$, $\beta_2 = \beta_2^f$ and $\beta_3 = \beta_3^f$ apply (Meese and Rogoff, 1983). However, any analysis in which the coefficients are restricted to be equal for each variable typically results in biased coefficients (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 (1982). A model that explicitly takes account risk premia into account is the portfolio balance model (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 in both monetary and portfolio

¹⁰ However, note that using the cumulated bilateral trade balance as a proxy for net foreign assets covers only a part of the current account. Besides the transfers, income and trade in services are excluded. Since returns on capital dominate the income variable the latter depends predominantly on returns such as interest rates which are included. Since trade in services was a minor issue over large parts of the sample it is reasonable to exclude it in our study.

¹¹ T denotes tradeables and NT non tradeables.

models (Frankel, 1983). As a consequence, we remove the restrictions of parameter equality of the interest rate differential and the inflation rate differential in equations (4), (5), (6), and (7).

Thus, we start our analysis as unrestrictive as possible and bear in mind dynamics stemming from both the portfolio balance approach and the monetary approach. Finally, we have four different models which all rely on the baseline specification of the unrestricted RID model exchange rate determination equation in equation (4).

2.2 Long-run analysis with time-varying coefficients

Wolff (1987) mentions three reasons why a time-varying coefficient model should be superior to fix coefficients 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 (see also Leventakis, 1987). Another reason is the famous Lucas critique: coefficients change if an anticipated change in the policy regime occurs. The third argument is related to the long-run real exchange rate. The monetary model assumes that 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. Because we explicitly account for changes in the real exchange rate the latter issue deserves less attention in our analysis with respect to the choice of the estimation technique.

A reason for choosing time-varying coefficient models can also be derived from different theories. In inter-temporal new open economy macroeconomic (NOEM) models (Obstfeld and Rogoff, 1995), money demand does not depend on income, but on real consumption. If we proxy real consumption by real income, a change in the average rate of consumption results in a change in the elasticity of income in the exchange rate equation. Thus, if consumption shares do vary which is, for instance, true for the US the exchange rate determination equation is thus also time-varying.

As argued by Wilson (1979), an anticipated policy change, i.e. an expansionary monetary policy, can generate dynamics which are different from that stemming 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.

According to the results gained by Sarno, Valente and Wohar (2004) or de Grauwe and Vansteenkiste (2007), the adjustment of exchange rates towards the long-run equilibrium relationship does not appear to be time-invariant. However, we expect that adjustment differs from period to period, at least over a long span of data. An adjustment towards the long-run equilibrium relationship can occur because the exchange rate predominantly reacts on the fundamentals or because, vice versa, the fundamentals react to changes in exchange rates. In the latter case, it is possible that the exchange rate does not adjust in sub-periods. Consequently, the adjustment coefficient has the potential to differ between sub-periods.

Siklos and Granger (1997) develop a framework well-suited to analyze these issues in the necessary detail. 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. This concept of regime-sensitive cointegration can be combined with a time-varying coefficient approach as follows. Let X_t^1 , X_t^2 and Y_t be different processes where

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

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

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

The variables 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.

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

with

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

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

Imposing cointegration on 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:

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

Equation (13) is a cointegration relationship if $\phi_t - \beta_t^1 \phi_{2t}^1 - \beta_t^2 \phi_{2t}^2$ 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, these parameters can be present or absent in sub-periods. This result is independent of 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 one of the stochastic trends in equation (8) is currently absent, the corresponding X_t^k variable does not enter the cointegration vector and the cointegration vector only contains two elements. Cointegration is continuously present over the whole period of observation while merely the composition of the cointegration vector is changing. If a system has at least one continuous common stochastic trend, Y_t continuously cointegrates with X_t^k only under the condition that $\phi_t - \beta_t^1 \phi_{2t}^1 - \beta_t^2 \phi_{2t}^2$ is zero. With $ect_t = \varepsilon_t^y - \beta_t^1 \varepsilon_t^{x1} - \beta_t^2 \varepsilon_t^{x2}$, the error-correction term therefore turns out to be

$$ect_t = Y_t - \beta_t^1 X_t^1 - \beta_t^2 X_t^2, \quad (14)$$

for which the error-correction presentation results as follows

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

where η_t is a i.i.d. variable which follows a normal distribution with zero mean. In addition to a time-varying cointegration vector, we allow the causality between the variables to change during the period of observation. This means that the dimension 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) does not only change its magnitude, it can also be zero if Y_t does not adjust at all to the long-run relationship.

In a long-run relationship analysis we thus are potentially confronted simultaneously with switch on and off cointegration, a changing cointegration vector and the adjustment

process. The difficulty with our estimations then is to cope with potential overlaps of these phenomena. Hence, our approach takes account of different regimes. Hence, it is able to distinguish between cases in which the cointegration relationship is switched on and those in which different adjustments are present.

For a multivariate case we consider the term

$$Y_t = \mu_t + \beta_t X_t + \varepsilon_t \quad (16)$$

with

$$X_t = [X_t^1, \dots, X_t^K] \text{ for } n = 1, \dots, K, \quad (17)$$

where K represents the maximum number of explanatory variables.¹² The matrix X_t has the dimension $(K \times 1)$ and β_t the dimension $(1 \times K)$. In our empirical analysis, we put the following models under closer scrutiny:

Model one:

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

Model two:

$$Y_t = [s_t], \quad X_t = [m, y, i, \pi, m^f, y^f, i^f, \pi^f, BCTB] \quad (19)$$

Model three:

$$Y_t = [s_t], \quad X_t = [m, y, i, \pi, m^f, y^f, i^f, \pi^f, \Delta CTB, \Delta CTB^f] \quad (20)$$

Model four:

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

¹² μ_t is a regime-dependent constant term. The variable ε_t represents an error term.

3. Modeling structural changes and estimating cointegrating relations - methodological issues

3.1 Testing for multiple structural changes

In general, *two frameworks* for tests for structural change can be distinguished. The first framework consists of 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 has to be rejected, meaning a structural change occurs at the corresponding point in time (Zeileis et al, 2003). The classical and the OLS based CUSUM test and the fluctuation test of Nyblom (1989) are well-known examples of these methods. These 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) which heavily relies upon Nyblom can be applied. While these procedures have the advantage of not assuming a particular pattern of deviation from the null hypothesis they can either only identify a single break or show general instability.

The second framework to test for structural changes is to compare the OLS residuals from regressions for different subsamples. This can be done, for example, by applying the F-statistics or the Chow test. In this paper, we adopt an extension of the latter case developed by Bai and Perron (1998, 2003). Their basic idea is to choose breakpoints such that the sum of squared residuals for all observations is minimized.

As a starting point, consider a multiple linear regression with m breakpoints and $m+1$ regimes

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

for $j=1, \dots, m+1$ with the convention that $T_0 = 0$ and $T_{m+1} = T$. y_t is the dependent variable, x_t' and z_t' denominate the regressors and κ and δ are the coefficient vectors. Note that only δ varies over time while κ is constant.

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 breakpoints (T_1, \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:

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 is imposed. In the framework of this paper, breaks are allowed to occur every 12 months.

$$(\hat{T}_1, \dots, \hat{T}_m) = \arg \min_{T_1, \dots, T_m} S_T(T_1, \dots, T_m). \quad (23)$$

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 breakpoints is also obtained by calculating the sum of squared residuals. To select the dimension of the model we apply the Bayesian Information Criterion (BIC) which according to Bai and Perron (2001) works well in most cases when breaks are present. After calculating the tests for all possible breakpoints the sequence $(\hat{T}_1, \dots, \hat{T}_m)$ is selected as the configuration at which the BIC achieves its minimum. Carrioni-Silvestre and Sanso (2006) show that this approach yields a consistent estimate of the break fraction. The breakpoints obtained in this fashion are a local minimum of the sum of squared residuals given the number of breakpoints but not necessary a global minimum.

It is important to note that the procedure of Bai and Perron has originally been developed for the case of stationary variables ($I(0)$). Nevertheless, it can as well be applied to non-stationary variables which are integrated of order one ($I(1)$). For instance, Siklos and Granger (1997) use this methodology to identify structural breaks in the interest parity equation between the United States and Canada in the context of regime-sensitive cointegration. In addition, Zumaquero and Urrea (2002) point out that the break estimator is consistent also in the non-stationary case. Using disaggregated price indexes for seven countries, they test for structural breaks in the coefficients of cointegrating relations which represent absolute and relative purchasing power parity. They also examine instabilities in the adjustment behaviour of price ratios and exchange rates. Finally, Perron and Kejriwal (2008) demonstrate 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 correct in both cases.

To check our results for robustness, we also apply the CUSUM test combined with Andrews and Ploberg (1994) in a similar way as Goldberg and Frydman (2001) to detect

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 breakpoints is then selected by comparing the F-values described above for the different numbers of breakpoints 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 "1 vs. 1+1" break tests. For details see Bai and Perron (1998, 2003).

¹⁵ This is only true if, as in our case, the intercept is allowed to change across segments.

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

possible breakpoints. However, with no considerable differences arising from the results, we proceed using the breakpoints obtained by the Bai and Perron methodology.

3.2 Estimating cointegrating relations with single equations

After identifying the breakpoints we now turn to the issue of correct estimation. As Bai and Perron's 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).¹⁷ 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.

For our purposes, the fully modified (FM) estimator is the most suitable method. In contrast to traditional single equation formulas it considers endogenous 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 this procedure is reliable in the case of full rank or cointegrated $I(1)$ regressors¹⁸ 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.¹⁹

The root idea of this concept is to estimate cointegrating relations directly by correcting traditional OLS with regard to endogeneity and serial correlation (Phillips, 1995). Let z_t denote 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

¹⁷ For a review of the different estimation methods of estimating cointegrating relationships see Hargreaves (1994), Phillips and Loretan (1991) and Caporale and Pittis (1999).

¹⁸ Note that the direction of cointegration does not need to be known. Regressors containing a deterministic trend are also allowed.

¹⁹ Furthermore, also Phillips and Hansen (1990), Hargreaves (1994) and Cappuccio and Lubian (2001) report good finite sample properties of the FM-OLS estimator.

regressors. u_t represents an n -vector stationary time series. Both vectors can be partitioned as follows:

$$z_t = \begin{bmatrix} y_t \\ x_{1t} \\ x_{2t} \end{bmatrix} \quad u_t = \begin{bmatrix} u_{1t} \\ u_{2t} \\ u_{3t} \end{bmatrix} \quad (24)$$

The data generating process of y_t is represented by the following cointegrated relation

$$y_t = \beta x_{1t} + u_{1t}. \quad (25)$$

The vectors of the regressors are specified as follows

$$\Delta x_{1t} = u_{2t}. \quad (26)$$

$$x_{2t} = u_{3t}. \quad (27)$$

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 non-stationary components, this transformation reduces asymptotically to the ideal correction while the differenced stationary components vanish asymptotically. Such a correction does not have any effect on the sub-vectors of x_t where serial correlation or endogeneity are not present.²⁰ A further advantage is that we do not have to account for cointegration between the x_{1t} regressors 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 auto-covariances (Hargreaves, 1994).

$$\Omega = E(u_t u_t') + \sum_{k=2}^{\infty} E(u_0 u_k') + \sum_{k=2}^{\infty} E(u_k u_0'). \quad (28)$$

$$\Omega = \Sigma + \lambda + \lambda'. \quad (29)$$

We define Δ as

²⁰ Without serial correlation or endogeneity the FM-OLS estimator is identical to the OLS estimator.

$$\Delta = \Sigma + \lambda. \quad (30)$$

Estimation of these covariance parameters can be achieved by using the pre-whitened kernel estimator suggested by Andrews and Monahan (1992).²¹ The endogeneity correction then has the form

$$y_t^* = y_t - \hat{\Omega}_{0x} \hat{\Omega}_{xx}^{-1} \Delta X_t. \quad (31)$$

The above correction is employed to account for endogeneities in the regressors x_{0t} linked with 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_{0t} 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

$$\hat{\Delta}_{0x} = \hat{\Omega}_{00} - \hat{\Omega}_{0x} \hat{\Omega}_{xx}^{-1} \hat{\Omega}_{x0}. \quad (32)$$

The correction is then given by

$$\hat{\Delta}_{0x}^* = \hat{\Delta}_{0x} - \hat{\Omega}_{0x} \hat{\Omega}_{xx}^{-1} \hat{\Delta}_{xx}. \quad (33)$$

Combining both corrections the formula for the fully modified estimator is²²

$$\hat{\beta}^* = (Y^* X - T \hat{\Delta}_{0x}^*) (X' X)^{-1} \quad (34)$$

3.3 Regime shifts in Cointegration models

To apply the FM-OLS estimator in a model with structural changes we proceed in a similar way as Hansen (2003) does in the Johansen framework by allowing the parameters to change their values at the breakpoints.²³

²¹ 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).

²² The traditional OLS estimator is given by $\hat{\beta} = Y' X (X' X)^{-1}$

We rewrite equation (22) with $\tau (t)$ as a constant

$$y_t = \tau(t) + x_t' \kappa(t) + z_t' \delta_j(t) + u_t \quad (35)$$

The piecewise constant time-varying parameters are given by

$$\delta_j(t) = \delta_1 1_{1t} + \dots \delta_m 1_{mt} \quad (36)$$

$$\kappa_j(t) = \kappa_1 1_{1t} + \dots \kappa_m 1_{mt} \quad (37)$$

$$\tau(t) = \tau_1 1_{1t} + \dots \tau_m 1_{mt} \quad (38)$$

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

$$1_{mt} = 1(T_{j-1} + 1 < t < T_j), J=1, \dots, m \quad (39)$$

with the convention that $T_0 = 0$ and $T_m = T$. Defining dummies according to the indicator function ensures that we are able to obtain 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 consumer price index (CPI). Furthermore, we use the overall trade balance as an approximation of the cumulated current account. 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 the case of our analysis a separate cointegration analysis (not reported) shows that the US dollar denominated balance adjusts to the euro denominated one. Thus, we choose the euro configuration. For the short-term interest rates we use money market rates with a maturity of three months. Exchange rates, money supply and real income are expressed in logarithms. All series are seasonally adjusted and are taken from International Financial Statistics of the International Monetary Fund.

²³ We corroborated our results with a related approach introduced by Gregory and Hansen (1996). They model the changes in the intercept and the slope coefficients relative to the first subperiod as a benchmark, running from 0 to T_1 . The base model is then written as $y_t = \tau_1 + \tau(t) + x_t' \kappa_1 + x_t' \kappa(t) + z_t' \delta_1 + z_t' \delta_j(t) + u_t$.

In strong contrast to other studies investigating the euro exchange rate, we rely on the Deutschmark and the fundamentals of Germany before the introduction of the euro. The reason is that we are interested in market rates which could be contrasted by using weighted ECU-Data. In a sense, the Deutschmark has been a predecessor of the euro as it had a similar importance on the foreign exchange markets as its successor. One reason was the big influence of the German Bundesbank (Fратиanni and von Hagen, 1990). We therefore use a time series which contains the German values until December 1998 and, from then on, the values of the euro area. Consequently, the Deutschmark / US dollar exchange rate is converted by the official Deutschmark / 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 euro area data. Since we deal with structural break models in the empirical section, we do not see any problems with our proceeding. The reason is that if a break due to data adjustment were important, the Bai-Perron test would signify a break around January 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 non-stationary variable to adjust, nor is a stationary relationship between $I(1)$ and $I(2)$ variables possible. Furthermore, inferences in a model with $I(2)$ variables are 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 instance we test for stationarity in the levels. Differences are taken and tested again 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 $I(1)$. In the case of the cumulated overall trade the results of the tests suggest that the balance is integrated of order two. Therefore, we decide to work with differences for the US and the euro area series. This can be done without changing the underlying economic theory. The results of the tests are presented in Table 1. According to the results, all variables can be considered as being integrated of order one although, in some cases, the evidence is mixed.

- Table 1 about here -

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 rejected for the change in the trade balance of the euro area according to the DF-GLS and the Phillips-Perron test. However, since the other tests indicate contrary results for these series we treat them as $I(1)$.

4.3 Empirical results

4.3.1 Estimation of the breakpoints

The breakpoints we are able to identify by applying the Bai and Perron methodology are presented in Table 2. Obviously, breaks occur quite frequently. Hence, we conclude that there is no stable and unique long-run equilibrium relationship among fundamentals and exchange rates since the breakdown of Bretton Woods I. Another result is that, despite a couple of differences, also some significant similarities between the various configurations emerge. For instance, the number of breakpoints always lies between eight and ten even though we allow for a shift every twelve months. Furthermore, the dates of breakpoints for the different models are located closely together. An encouraging result is that, in many cases, major economic or political developments are able to deliver “good” explanations for instabilities.

- Table 2 about here -

The breaks in 1976 and 1977 (row 2 of Table 2) can clearly be addressed with an eye on the macroeconomic turbulences arising from the oil price shocks and worldwide recession. Furthermore, instabilities often occur within the epoch of the so-called pseudo-monetarism policy of the FED within 1979 and 1982 (Timberlake, 1993) or at the end of the rise of the US dollar during the mid 1980s. From this point of view, we feel legitimized to explain the breaks of model 1 and 3 in a textbook-style fashion by the beginning of the monetary “experiment” (row 3 of Table 2). In addition, the regular interventions by the treasury were stopped as had been announced in April 1981 although until 1985 infrequent interventions occurred (c.f. Destler and Henning, 1989). In order to support the real economy, the federal funds rate started to fall in mid 1981. This date coincides with breakpoints in each model (row 4 of Table 2).

The next breakpoint located around October 1988 (row 7 of Table 2) can also well be traced back to a specific stance of monetary policy. In 1988, the monetary policy stance on both sides of the Atlantic, i.e. of both the US Fed and Germany’s Bundesbank, became more restrictive. Besides the usual monetary policy suspects, the election of George Bush senior

and the G-7 summit in Berlin²⁴ offer further popular explanations.

For each model, breakpoints are identified in February 1992, shortly after the German reunification (row 8 of Table 2). The following instability in 1992 and 1993 (row 9 of Table 2) is usually attributed to the crisis of the European Monetary System. At this time, also significant changes in US and German monetary policies have to be taken into account.

Within a comparatively stable period until the end of the 1990s, the only instability, in 1997 (row 11 of Table 2), is said to be caused by the Asian currency crisis and/or the worsening of the US trade balance which had started in 1996. Afterwards, breaks are reported by the Bai and Perron procedure for model one, two and three in 2000 (row 13 of Table 2) and for each model in November 2004 (row 14 of Table 2). In mid-2000, the American economy started to slow down with the American stock market crashing. Interestingly, the last break coincides exactly with an event which saw the short-term interest rates of the euro area declining below the level of US interest rates. Of course, as far as the dating of breakpoints and their economic interpretation are concerned, we preferred to follow quite standard interpretations. Moreover, one should also not forget that many other important developments are not reflected by breaks. However, in all these cases the interest rate differential seems to play an important role. As becomes obvious after a visual inspection of Figure 1, many breaks correspond to and are thus potentially triggered by changes in the trend of the interest rate differential or to changes in its sign.

- Figure 1 about here -

4.3.2 Interpretation of the time varying coefficients

Moving one step further, we proceed by estimating the cointegration vector via FM-OLS using the obtained break dates. Table 3-6 contain the results for the specified models. Since 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 configuration 1 (Table 3) just for comparison. In order to take account of the different model specifications proposed by Hooper and Morton (1982), we draw on model 2 and model 3 to distinguish between the bilateral net foreign asset position and the overall net foreign (nfa) asset positions of each country (in our case, the changes in the nfas). A comparison of model 2 and 3 with 4 helps us in separating real effects as the latter case doesn't account for changes in the trade balances.

- Table 3-6 about here -

²⁴ In contrast to previous meetings, the participants of the Berlin meeting did not publically claim that fluctuations in the dollar were unwanted.

Models 1, 2 and 3 are broadly consistent with the real interest rate model (Equation 4) in the first sub-period after our period of observation starts (row 1 of Tables 3-5). From this point of view, our empirical findings clearly corroborate the findings in the literature concerning the early period after the breakdown of Bretton Woods I (for an early overview see, for instance, Isard, 1987). Only in the case of model 4 does the German inflation expectations variable enter the regression equation with an incorrect sign of its estimated coefficient (row 1, column 5 of Table 6). While the overall change in the overall net foreign asset position (nfa) of the euro area/ of Germany in model 4 is not significant, the same variable turns out to be significant at the 1% level in the US case (row 1, columns 10 and 11 of Table 6). Its negative sign indicates that risk considerations seem to be important. During this period, a worsening of the US trade account is linked to a depreciation of the US dollar. It is important to note that the US money supply seems to be strongly linked to the exchange rate. During this period both variables appear to share common trends. (row 1, column 6 of Tables 3, 5 and 6).

From 1977:05 till 1979:12 many coefficients of model 3 show signs which are not consistent with standard theory (row 2 of Table 5). The estimated coefficients of both the German money supply and the German inflation expectations turn out to be highly significant with a negative sign (row 2, column 2 and 5 of Table 5). When either the relative price of tradeables (row 2, column 10 and 11 of Table 6). or the bilateral nfa (row 2, column 10 of Table 4). is taken into account in model 2 and 4, their coefficients display the correct sign and the significance of the money supply and the inflation rates disappears. One can think of several reasons for that pattern. On the one hand, 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 took place exactly in this period. It becomes obvious that real shocks have an impact on the exchange rate and let the impact of nominal factors shocks vanish.

The pattern of the estimation results for the sample period from 1979:12 till 1981:06 in model 3 are again broadly consistent with the theory (row 3 of Table 5). Despite the fact that the above mentioned episode of the "monetary experiment" initiated by the US Fed falls in this period, it becomes obvious that the coefficients for US money supply and inflation rates are in line with the theory, i.e. impacts of US monetary variables determine the exchange rate. This is particularly true for model 1 and 3 (row 3, columns 6 and 9 of Tables 3 and 5). The only deviation from the real interest rate model (equation 4) is that the German short-term interest rates enter with a positive sign which indicates that the opportunity costs of holding money are important in the short run.

Between 1981 and the end of 1984 (Model 1) respectively the beginning of 1985 (Model 4) the estimated coefficients of US money supply and the US real income variable show signs which are not consistent with standard theory. (row 3, columns 2 and 3 of Tables

3 and 6). Following the broad picture conveyed in Figure 2 which displays the time series of the macroeconomic indicators in the United States from 1973 until 2007, we attribute this pattern to the deepening recession in the US economy. However, if relative prices of tradeables are included, this yields signs of the estimated coefficients which are broadly consistent with the underlying theory of section 2.1 (row 3, columns 10 and 11 of Table 6). The estimated coefficient of the bilateral cumulated trade balance (row 3, column 10 of Table 4) reveals a positive sign which means that an increase of German claims on US assets coincides with a depreciation of the DM vis-à-vis the US dollar. As the US dollar appreciated strongly against major currencies during this period, such a correlation can be explained by an overshooting process as a result of anticipated monetary shocks: an announced monetary expansion causes a currency appreciation while the money stock widens. The money inflow generates current account deficits. This linkage is reflected by the positive sign of the estimated coefficient of the bilateral nfa. The negative sign of the estimated coefficients of German monetary variables can well be traced back to these events. The German central bank turned towards a looser monetary policy and the inflation rates slumped at the same time.

- Figure 2 about here -

The following period (1984:07 to 1988:10 for model 1 and 1984:07 to 1988:08 for model 3, 1985:03 to 1988:10 for models 2 and 4) is characterized by interventions which should have weakened the US dollar.²⁵ In all models, the estimated coefficients of inflation expectations in Germany are highly significant while the estimates of US real income shows mostly an incorrect sign based upon standard theory (row 4, columns 3 and 5 of Tables 3-6). This is largely due to the interventions occurred. In the next period, which starts in 1988:10 (except for model 3, which starts in 1988:08) and ends in 1991:02, all signs are broadly consistent with the theory (row 5, of Tables 3-6). The results indicate that liquidity effects are important for Germany. In the aftermath of the economic recovery inflation improved. As a consequence, the US and German monetary policy reacted, whereas the Fed raised interest rates first. The strong impact of money supply and inflation rate expectations can be seen in all models. 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 (row 6, columns 5, 10 and 11 of Tables 4 and 5). Besides German reunification, which caused a jump in the German money stock, the interest rate differential between the US and Germany changed its sign (see Figure 1). The signs of the estimated coefficients of the interest rate variables in all models support the view that

²⁵ The standard interpretation is that the Plaza agreement should have depressed the US dollar while the Louvre accord is said to have terminated the inclination of the US dollar towards depreciation.

Germany's advantage in interest rates initiated an appreciation of the Deutschmark against the US dollar.

After the crisis of the European Monetary 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. 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 the coefficients and their signs to be similar. Obviously, the inclusion of either the bilateral net foreign asset position, overall net foreign position or relative prices of tradeables changes the results considerably (row 7, of Tables 3 to 6). For model 1, 2 and 3 a further regime starts during 2000 (in 2000:01 for models 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. Regarding the estimated coefficients, the period between the end of 1993 and the beginning of 2000 is absolutely incompatible with standard theory.

The only analogy in fundamentals can be observed with respect to inflation rate expectations and US short-term interest rates. Both seem to be of equal importance. The common starting date of this period can be implicated in the establishing recession in Germany (see Figure 3). The breakpoint occurs exactly when the recession achieved its peak. At the same time German interest rates fell, which initiated a turnaround in the interest rate differential. In model 1, the next break occurs at a point in time at which the German interest rate differential became negative and in model 3 when the upward tendency stopped. This can be an explanation because the sign of the estimated coefficients of US interest rates changed from the preceding to the next regime. A reason for this additional break in model 3 can also be attributed to the use of the changes in the overall net foreign asset position. The changes of overall net foreign asset position 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 dominates the effects.

- Figure 3 about here -

The breakpoints for models 1, 2 and 3 occur when both the US and the euro area economy²⁶ started to slow down with falling US inflation rates and short-term interest rates (see Figures 2 and 3). Again, the interest rate differentials from the euro area to US started to narrow. The first years of the euro also yield results for the estimated coefficients of both

²⁶ Note that, at this stage of analysis, we switch to the use of euro area data.

the euro area and the US money supply and real income, both of which show signs that oppose standard theory. Nevertheless, the estimated coefficients of inflation rates in all models (except the US one in models 2 and 4) and the estimated coefficients of relative prices of tradeables in model 4 display the correct signs. From this point of view, inflation rate expectations and real effects had an important impact during this period. The last regime, which is the same in all four models, seems to be characterized by overshooting (Frankel (1979)) because the estimated coefficients of interest rates, broadly speaking, reveal the corresponding sign. This is in line with the interest rate differential. The euro area interest rates exceed the US interest rates when the break is located. The change in the signs of the coefficient from the last regime to the next supports this finding.

To sum up the findings, the Deutschmark/euro always appreciates against the US dollar when German respectively euro area interest rates are higher than interest rates in the USA. This tendency is driven by both interest rates. However, in all other periods the liquidity effect seems to dominate for the Germany/euro area, whereas no clear picture emerges for the USA. 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 their estimated coefficients are significant the sign changes frequently. Nevertheless, there is only one period in which the estimated coefficient of the change in overall nfa has the same signs, namely the one ranging from 1997:06 until 2000:01.

In model 4 all coefficients of the US foreign prices have the same sign, i.e. an increase in the US relative price of tradeables 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 estimated coefficient displays the wrong 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 that, based upon the results of model 4, 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 the 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 the period from 2004:1 to 2007:12 inflation expectations concerning the USA became more important and as a consequence the relative price of tradeables is less important.

Finally, we can conclude that the relationship between exchange rates and fundamentals over a period of at least one and a half years 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 rates and fundamentals differs in each period.

4.3.3 Analysis of the error-correction term

In the last part of our analysis we examine whether the estimated relationship can be interpreted as a cointegration vector. As a first step we apply unit root tests to the error series obtained from the FM-OLS estimation following the idea of residual based cointegration tests. In doing so, we have to use the critical values for cointegration analyses which take account for the number of estimated parameter. Because of the large number of parameters used in our estimation we cannot rely on the standard critical values provided by the literature. For this reason, we ran a Monte-Carlo simulation with 10.000 repetitions in order to obtain critical values for our models.²⁷ According to the results of the DF-GLS and the Phillips-Perron Test which are reported in Table 7, the error term resulting from the step-wise relationships should be considered as stationary which gives clear evidence in favour of a long-run cointegrating relationship between exchange rates and fundamentals. This is also an indication for an error-correcting behaviour, meaning that the exchange rate endogenously adjusts to disequilibria.

- Table 7 about here -

An interesting question is whether this error-correction mechanism is also subject to structural changes. To tackle this question we apply the Bai and Perron test once again but in the following without imposing any restriction on the minimum distance between two breaks. The results which are summarized in Table 8 show that we are able to identify four breakpoints for model 2 and three nearly equal breakpoints for the other models.

- Table 8 about here -

Furthermore, a regression of the change in the exchange rate on the error term shows that the deviation of the exchange rate from its equilibrium determined by the cointegrating relation is significant and, as expected from theory, enters with a negative coefficient. The corresponding results which are summarized in the Tables 9 show that this is always true except for the first period of model 2 which only lasts 8 months. As can be seen by looking at the estimated coefficients, the constant term is mainly responsible for the

²⁷ To be more precise, we construct the data generating process for each variable. Each process is constructed as an independent random walk. In addition, we take account for the breaks obtained by each model. Consequently, the null hypothesis is no cointegration, meaning that we obtain a series for the error term that contains a unit root for each model. The critical values can then be drawn from the realized distribution. However, this methodology cannot be applied to the KPSS test which assumes stationarity under the null. In this case, we would need to know the exact specification of the cointegration relationship under the consideration of our breaks to obtain relevant critical values. We therefore decided to leave out the KPSS test and to rely on the DF-GLS and the Phillips-Perron Test.

breaks found up to the end of the 1980s because the regimes perfectly coincide with long-swings in the exchange rates. Thus, only in the last period there is some evidence that the adjustment speed has shrunk.

- Table 9 about here -

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 breakpoint 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. Conclusions

In this paper, we have examined the long-run relationship between the US dollar/euro exchange rate and fundamentals with respect to structural breaks in the coefficients. We show that fundamentals are important in each sub-period but that their impact differs significantly depending on different regimes. With respect to this issue we draw some major conclusions.

One result we come up with is that there are no perseverative regimes, implying that either the empirical realisations of the estimated coefficient for the same fundamentals or their significance values differ. Insofar as efficient forex market intervention presupposes the exact knowledge of the dollar/euro equilibrium exchange rate, this makes exchange rate targeting a technically demanding exercise because it has to deal with a moving target. Moreover, our results contradict the view that fundamentals only matter during single periods while having no explanatory content within other regimes. Hence, we would like to argue that our specification beats the pure random walk hypothesis in the Meese and Rogoff (1983) sense. Goldberg and Frydman (2001) offer a possible explanation of our findings. In their view, market participants change the theories respectively the fundamentals they use to forecast exchange rate movements. Those changes in turn influence the paths of the exchange rate (Christopoulos and León-Ledesma, 2009). Furthermore they could well be explained by the specific economic events we address to explain our findings in chapter 4.

In technical terms, we were able to establish the existence of cointegrating relations by testing the respective error terms for stationarity. Moreover, the dollar/euro exchange rate significantly adjusts to deviations from the step-wise linear relationships in all cases.

Altogether, modelling the dollar/euro exchange rates in a linear fashion appears to be inadequate in many instances. Thus, we feel legitimized to claim that the poor empirical record of some standard monetary exchange rate models can be attributed to, among other

factors, the assumption of regression coefficients which do not change over time. Another result is that, in several instances, specific economic developments can well be identified and addressed to explain the date of the breaks. The same is true concerning the character of estimated relationships between the reported fundamentals and the exchange rate for the different periods.

The topic addressed by us surely needs further attention. While our focus has been on the exchange rate, an analogous study could also be conducted for the extensive evidence of parameter instability seen in other (forward looking) macroeconomic and financial data. Separate from the interesting question of what accounts for the time-varying relationship between exchange rates and fundamentals, there is also the question of what its implications are (Bacchetta and van Wincoop, 2009). We leave the task of corroborating our results for other currency pairs or other model configurations to further research.

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Tables and Figures

Table 1 - Unit root tests

Variable	Levels				First Differences				
	PP		DF-GLS		PP		DF-GLS		KPSS
	test statistic ^a	lags	test statistic ^b	test statistic ^c	test statistic ^a	lags	test statistic ^b	test statistic ^c	
<i>EUR / USD</i>	-1.317	2	-0.437	2.690**	-16.66**	0	-1.485	0.084	
<i>m</i> ^{EMU}	-1.662	0	-1.691	1.008*	-21.800*	0	-19.335*	0.123	
<i>y</i> ^{EMU}	-3.36	15	-2.693	0.182**	-31.059*	0	-25.513*	0.049	
<i>i</i> _s ^{EMU}	-1.97	0	-1.154	1.840**	-19.86**	0	-17.069*	0.074	
<i>π</i> ^{EMU}	-2.594	12	-0.651	2.012**	-17.32**	0	-7.782**	0.11	
<i>ΔCTB</i> ^{EMU}	-4.048*	0	-4.643*	0.566*	-31.772*	0	-30.161*	0.062	
<i>m</i> ^{US}	-0.027	8	-0.669	1.543*	-15.202*	16	-2.121**	1.696*	
<i>y</i> ^{US}	-1.839	0	-1.253	0.489*	-15.268*	0	-3.335*	0.083	
<i>i</i> _s ^{US}	-1.899	12	-1.636	3.466*	-16.559*	0	-16.480*	0.456	
<i>π</i> ^{US}	-2.581	12	-0.373	3.551*	-13.701*	0	-13.606*	0.178	
<i>ΔCTB</i> ^{US}	-0.62	0	-0.974	1.336*	-28.596*	0	-16.376*	0.628**	
<i>BCTB</i>	-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 and 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. *m* denotes money supply, *y* real income, *i*_s short-term interest rates, *π* inflation rate expectations, *ΔCTB* the change in the cumulated trade balance and *BCTB* the bilateral cumulated trade balance. *EUR/USD* is the euro price of one unit US dollar. Sample period: 1975:01 to 2007:12.

Table 2 - Dating of breakpoints in monetary models of the exchange rate

	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: The reported breakpoints are obtained by applying the Bai and Perron (1997, 2003) methodology on the regression $Y_t = \mu(t) + \beta(t)X_t + \varepsilon_t$ for the different models described in section 2. Y_t contains the euro-US dollar exchange rate and X_t is a $K \times 1$ vector of K fundamentals of each model. Breaks within a horizon of 6 months are seen as comparable. Sample period: 1975:01 to 2007:12.

Table 3: Estimation results of model 1 (cointegrating relations)

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

Note: The results are obtained by regressing the exchange rate on fundamentals contained in model 1 (for a description of this model see section 2.2). The sub-periods are modelled by using indicator functions based on: $Y_t = \mu(t) + \beta(t)X_t + \varepsilon_t$. m denotes money supply, y real income, i_s short-term interest rates and π inflation rate expectations. P-values are in parentheses. * denotes statistical significance at the 10% level, ** at the 5% level and *** at the 1% level. Sample period: 1975:01 to 2007:12.

Table 4: Estimation results of model 2 (cointegrating relations)

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

Note: The results are obtained by regressing the exchange rate on fundamentals contained in model 2 (for a description of this model see section 2.2). The sub-periods are modelled by using indicator functions based on: $Y_t = \mu(t) + \beta(t)X_t + \varepsilon_t$. m denotes money supply, y real income, i_s short-term interest rates, π inflation rate expectations and $BCTB$ the bilateral cumulated trade balance. P-values are in parentheses. * denotes statistical significance at the 10% level, ** at the 5% level and *** at the 1% level. Sample period: 1975:01 to 2007:12.

Table 5: Estimation results of model 3 (cointegrating relations)

	μ	m^{EMU}	y^{EMU}	i_s^{EMU}	π^{EMU}	m^{US}	y^{US}	i_s^{US}	π^{US}	ΔCTB^{EMU}	ΔCTB^{US}
1975:01	9.109 (0.000)	*** -0.088 (0.552)	-0.447 ** (0.031)	-0.038 *** (0.000)	-0.475 *** (0.678)	-1.729 *** (0.000)	0.913 ** (0.016)	0.032 *** (0.000)	-2.883 *** (0.000)	0.008 (0.407)	-0.014 *** (0.004)
1977:05	-3.987 ** (0.049)	-1.536 *** (0.000)	-0.042 (0.810)	0.025 *** (0.000)	-3.207 ** (0.013)	0.536 (0.249)	0.325 (0.348)	-0.014 *** (0.000)	-0.635 (0.393)	0.000 (0.992)	0.004 (0.277)
1979:12	3.808 (0.567)	1.391 *** (0.000)	-1.008 *** (0.009)	0.016 *** (0.000)	7.810 *** (0.000)	-3.115 *** (0.000)	0.502 (0.217)	0.008 ** (0.014)	-8.485 *** (0.000)	0.026 *** (0.007)	0.019 *** (0.000)
1981:06	-8.491 *** (0.000)	-0.169 (0.460)	-0.441 *** (0.000)	0.047 *** (0.000)	-5.689 *** (0.000)	0.645 ** (0.018)	-0.375 * (0.085)	-0.002 (0.291)	-0.893 (0.114)	0.004 (0.458)	-0.006 (0.000)
1984:07	0.256 (0.876)	-0.098 (0.566)	0.309 ** (0.046)	0.081 *** (0.000)	7.791 *** (0.000)	-0.317 *** (0.006)	-2.002 *** (0.000)	-0.001 (0.773)	-1.140 ** (0.014)	-0.018 *** (0.000)	-0.001 (0.309)
1988:08	22.217 *** (0.000)	0.239 *** (0.001)	-0.291 (0.177)	0.011 ** (0.023)	0.351 (0.769)	-5.025 *** (0.000)	0.453 (0.214)	0.001 (0.874)	-2.383 *** (0.001)	-0.005 (0.188)	0.004 * (0.084)
1991:02	-10.491 *** (0.000)	-0.291 (0.403)	-0.162 (0.229)	-0.024 *** (0.000)	4.754 *** (0.000)	0.207 (0.487)	0.577 (0.208)	0.046 *** (0.000)	-0.298 (0.661)	-0.013 *** (0.000)	0.006 ** (0.010)
1993:12	-25.356 *** (0.000)	1.142 *** (0.000)	-0.034 (0.881)	0.072 *** (0.000)	6.895 *** (0.000)	0.601 (0.198)	0.852 * (0.063)	-0.051 *** (0.000)	-6.874 *** (0.000)	0.004 (0.148)	0.001 (0.593)
1997:06	-16.296 *** (0.000)	-1.003 *** (0.000)	0.902 *** (0.003)	-0.058 *** (0.000)	0.520 (0.657)	2.450 *** (0.000)	-1.541 *** (0.001)	0.053 ** (0.022)	-2.036 ** (0.282)	-0.007 *** (0.002)	-0.014 *** (0.000)
2000:01	6.302 *** (0.004)	-0.923 *** (0.000)	0.122 (0.664)	0.059 *** (0.000)	7.657 *** (0.000)	0.446 ** (0.028)	-2.729 *** (0.000)	0.013 ** (0.036)	-2.205 *** (0.000)	0.005 *** (0.000)	-0.004 *** (0.000)
2004:11	-13.322 ** (0.011)	0.733 *** (0.000)	-0.462 (0.260)	-0.099 *** (0.000)	-0.057 (0.974)	0.189 (0.732)	-0.200 (0.700)	0.016 * (0.059)	-1.493 ** (0.043)	-0.001 (0.455)	-0.002 (0.223)

Note: The results are obtained by regressing the exchange rate on fundamentals contained in model 3 (for a description of this model see section 2.2). The sub-periods are modelled by using indicator functions based on: $Y_t = \mu(t) + \beta(t)X_t + \varepsilon_t$. m denotes money supply, y real income, i_s short-term interest rates. π inflation rate expectations and ΔCTB the change in the cumulated trade balance. P-values are in parentheses. * denotes statistical significance at the 10% level, ** at the 5% level and *** at the 1% level. Sample period: 1975:01 to 2007:12.

Table 6: Estimation results of model 4 (cointegrating relations)

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

Note. The results are obtained by regressing the exchange rate on fundamentals contained in model 4 (for a description of this model see section 2.2). The sub-periods are modelled by using indicator functions based on: $Y_t = \mu(t) + \beta(t)X_t + \varepsilon_t$. m denotes money supply, y real income, i_s short-term interest rates, π inflation rate expectations and p a price index. T describes tradeable goods and NT non-tradeable goods. P-values are in parentheses. * denotes statistical significance at the 10% level, ** at the 5% level and *** at the 1% level. Sample period: 1975:01 to 2007:12.

Table 7: Unit root tests for the error terms

	PP	Critical values		lags	DF-GLS	Critical values	
	test statistic ^a	1% level	5% level		test statistic ^b	1% level	5% level
Model 1	-16.20***	-5.778642	-4.609888	2	-14.60***	-5.286206	-4.459672
Model 2	-14.67***	-5.667703	-4.530416	0	-14.54***	-5.212317	-4.358875
Model 3	-18.50***	-4.699536	-4.075982	0	-17.54***	-4.571048	-4.068024
Model 4	-15.06***	-7.604278	-5.113595	0	-14.34***	-6.530052	-4.753679

Note: * Statistical significance at the 5% level, ** at the 1% level. Both the PP test and the DF-GLS test assume that the series contains a unit root under the null. To obtain the relevant critical values we ran a simulation with a sample size of 10000 for each model. Sample period: 1975:01 to 2007:12.

Table 8: Comparison of breaks in the error-correction model

	Model 1	Model 2	Model 3	Model 4
	1975:01	1975:01 1975:09	1975:01	1975:01
	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

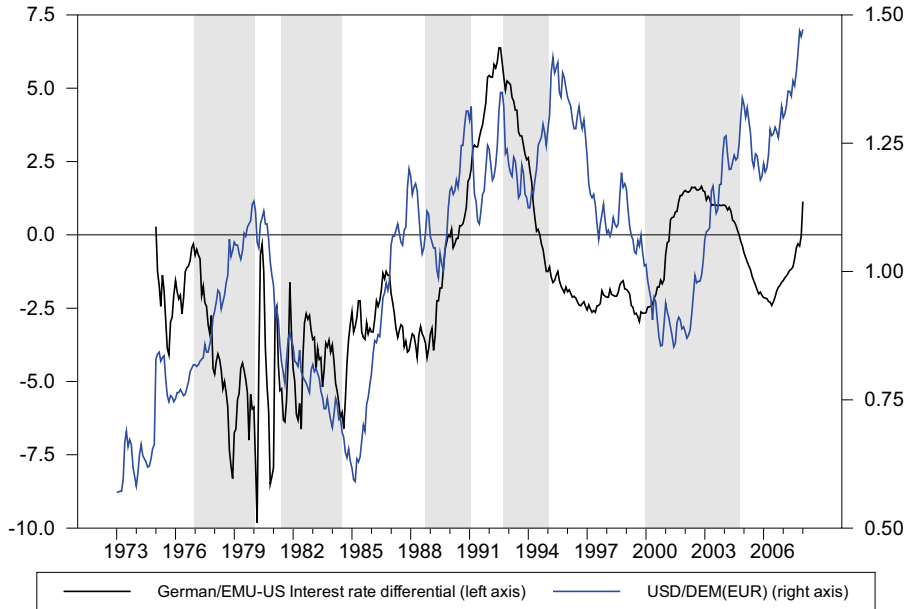
Note: The reported breakpoints are obtained by applying the Bai and Perron (1997, 2003) methodology on the regression $\Delta s_t = \mu(t) + \alpha(t)ect_{t-1} + \varepsilon_t$ for the different models. Breaks within a horizon of 6 months are seen as comparable. Sample period: 1975:01 to 2007:12.

Table 9: Error-correction estimations for each selected model

Sub-period	Model 1		Model 2		Model 3		Model 4	
	$\mu(t)$	$\alpha(t)$	$\mu(t)$	$\alpha(t)$	$\mu(t)$	$\alpha(t)$	$\mu(t)$	$\alpha(t)$
1975:01	-0.004 (0.104)	-0.340 ** (0.026)	0.016 ** (0.030)	-0.226 (0.554)	-0.004 (0.124)	-0.308 *** (0.003)	-0.005 (0.076)	-0.252 * (0.054)
1975:09			-0.024 (0.002)	-0.552 *** (0.000)				
1980:07	0.016 *** (0.003)	-0.558 *** (0.001)	-0.005 (0.499)	-0.706 *** (0.000)	0.016 *** (0.001)	-0.677 *** (0.000)	0.016 *** (0.000)	-0.734 *** (0.000)
1985:03	-0.023 *** (0.000)	-0.665 *** (0.000)	-0.042 *** (0.000)	-0.617 *** (0.006)	-0.022 *** (0.000)	-0.655 *** (0.000)	-0.022 *** (0.000)	-0.717 *** (0.000)
1987:02	0.003 (0.335)	-0.387 *** (0.000)	-0.017 *** (0.023)	-0.361 *** (0.000)	0.003 (0.376)	-0.383 *** (0.000)	0.004 (0.240)	-0.452 *** (0.000)

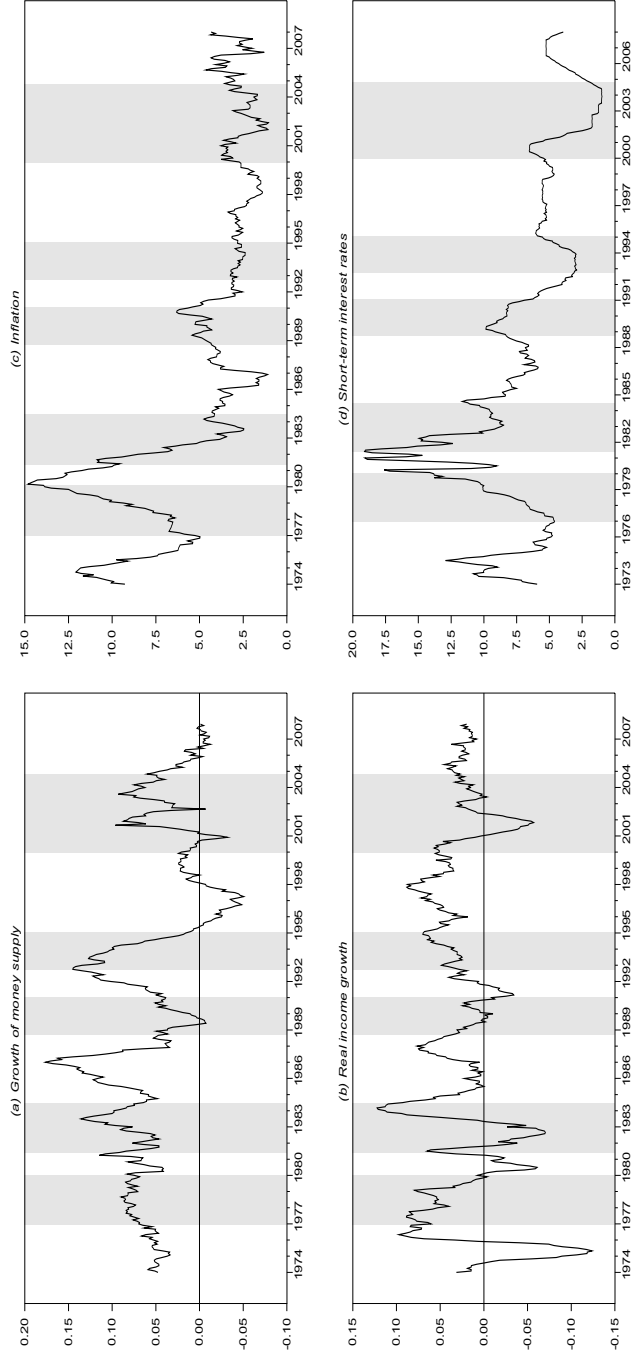
Note: The results are obtained by regressing the exchange rate in first differences on the one period lagged error term for each model (for a description of the models see section 2.2). The sub-periods are modelled by using indicator functions based on: $\Delta s_t = \mu(t) + \alpha(t)ect_{t-1} + \varepsilon_t$. P-values are in parentheses. * denotes statistical significance at the 10% level, ** at the 5% level and *** at the 1% level. Sample period: 1975:01 to 2007:12.

Figure 1: Interest rate differential and exchange rate - United States vis-à-vis the euro area



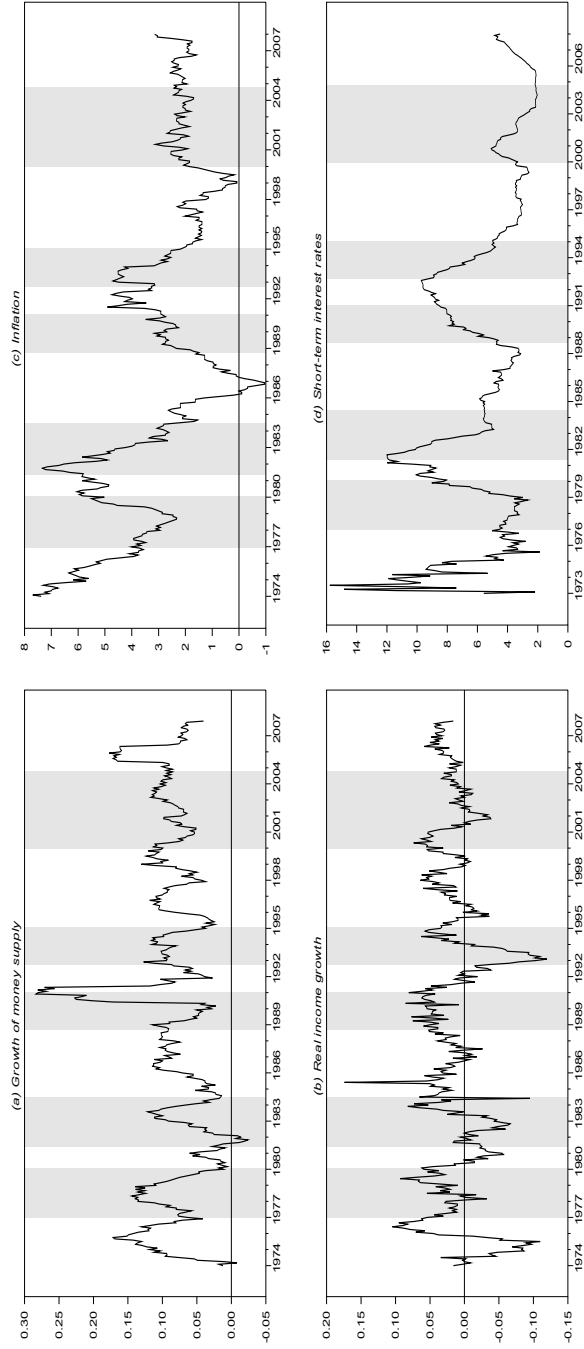
Note: The figure displays the development of the 3-month interest rate spread and the exchange rate between Germany (1975 until 1999) and the euro area (1999 until 2007), respectively, and the United States. Changes in the colour indicate a new regime. The regime classifications are based upon model 1.

Figure 2: Macroeconomic indicators in the United States from 1973 until 2007



Note: The graphs display the behaviour of four main macroeconomics indicators for the United States from 1975 until 2007. Changes in the colour indicate a new regime. The regime classification bases upon model 1.

Figure 3: Macroeconomic indicators in Germany and the euro area from 1973 until 2007



Note: The graphs display the behaviour of four main macroeconomics indicators for Germany (1975 until 1999) and the euro area (1999 until 2007). Changes in the colour indicate a new regime. The regime classification bases upon model 1.