Strong comovements of exchange rates: Theoretical and empirical cases when currencies become the same asset

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The aim of this paper is to detect periods in which two currencies can be classified as being the "same" asset. Two currencies can be treated as the same asset if their exchange rates vis-à-vis the same base currency are cointegrated with a cointegration vector that is consistent with the triangular arbitrage condition. In a first step, it is theoretically derived under which conditions, with respect to the process of the fundamentals, the exchange rates are cointegrated. The empirical results yield that periods of strong comovements of the US dollar and Pound sterling based upon the Euro prevail during the 1990s and periods of comovements of Euro and Pound sterling denominated in US dollar prevail since the introduction of the Euro. Furthermore, no long-run relationships can be discovered.
This paper gives four major innovations to the literature. It first shows under which conditions exchange rates can be bivariately cointegrated. Secondly, it uses the cross-rate identity to test for cointegration, i.e. deducing recursively. Thirdly, it applies the cointegration methodology within a triangular framework by detecting cointegration between exchange rates that are not only denominated in U.S. dollars. And lastly, it shows that comovements between two exchange rates exist in a narrower sense but only in short periods, whereas the economic variables which have caused the relationship are explored.

JEL classification: E44, F31, G15

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1 Introduction

In past decades much attention was directed to the explanation of bilateral exchange rates' movements. A big field of research surrounds the question: are future exchange rates predictable? A vast strand in literature is concerned with this topic and ultimately started with the seminal papers by Meese and Rogoff (1983a, 1983b), who tested fundamental theories of exchange rate determination like the monetary model and variations empirically and used them to make out-of-sample forecasts. Meese and Rogoff found that none of the fundamental models can outperform the results of a random walk forecast. Many subsequent papers broadly confirm the results of Meese and Rogoff by using fundamental models (Mark, 1995; Chinn and Meese, 1995 or Cheung and Chinn, 1998). On the other hand, some researchers can show predictability even at shorter horizons (MacDonald and Taylor, 1994; MacDonald and Marsh, 1997 or Cheung et al., 2005).[1] What all mentioned contributions have in common is that they investigate bilateral exchange rates and stress, more or less concretely, the linkage to fundamentals. Although most of the papers have investigated different exchange rates simultaneously long-run comovements of them are mostly neglected.

Another strand in the literature has been established following an argument Granger (1986) gives within the discussion fields of application in the cointegration literature. He argues that prices of speculative assets cannot be cointegrated because cointegration would contradict the weak form of market efficiency. Since a cointegration relationship is equivalent to the existence of an error correction term one price could be predicted by using another one. A second strand in the literature looking at long-run relationships of exchange rates applies the cointegration methodology to investigate for common stochastic trends. This is predominantly done to test for the stability within the European Monetary System before the introduction of the Euro in 1999 (e.g. Haug et al., 2000). The evidence of cointegration on the foreign exchange market is not clear. Some papers can reject the hypothesis of no cointegration (Baillie and Bollerslev, 1989; Septhon and Larsen, 1991; Norrbin, 1994; Lajaunie and Naka, 1997; Woo, 1999; Haug et al., 2000) while some others cannot (MacDonald and Taylor, 1989; Hakkio and Rush, 1989; Coleman, 1990; Copeland, 1991; Rapp and Sharma, 1999). The results are very sensitive to the period of observation and the model used (Sephton and Larsen, 1991; Jeon and Seo, 2003; Phengpis, 2006). Cointegration can mostly be considered between exchange rates that participate in the EMS before the introduction of the Euro. Kühl (2007) is the first to apply the methodology to the most traded exchange rates in terms of U.S. dollars since the introduction of the Euro. In this most recent period, for instance, bivariate cointegration exists between the most traded (in terms of market turnover) and the third most traded exchange rate, namely the Euro-U.S. dollar rate and the British pound-U.S. dollar rate.[2]

In one of the first contributions in the strand of literature concerned with market efficiency consideration, Hakkio and Rush (1989) verbally outline what is implicitly used in the convergence literature. They point out that "[i]f countries either explicitly (and rigidly) fix their exchange rates or implicitly link their economic policies [...], then their currencies are not different assets" (p. 87).

[1] MacDonald/Taylor (1994) apply a monetary model within an error correction framework and can outperform the random walk even at shorter horizons. MacDonald/Marsh (1997) use an expanded version of the purchasing power parity. Cheung et al. (2005) investigate not only monetary models but also test newer fundamental models that take, for example, productivity developments into account.

[2] In addition, cointegration can be detected for the British pound - Swedish Krona and Euro - Swedish Krona pairs denominated in U.S. dollar in sub-samples.

A similar argument is given by Baffes (1994). He writes "assets are different not only with respect to the physical properties [...] but also with respect to the set of fundamentals that determines their relative prices" (p. 277). Hence, this argument refers to factors that have an impact on more than just one exchange rate.

Dwyer and Wallace (1992), Baffes (1994), and Ferré and Hall (2002) show that two currencies can be treated as the same asset if their exchange rates vis-à-vis the same base currency are cointegrated with a cointegration vector that is consistent with the triangular arbitrage condition (Frenkel and Levich, 1975; Levich, 1985).[3] In Kühl (2007), this is true for the exchange rate pair Euro/ US-Dollar and British Pound/ US-Dollar. From this point of view, the Euro and the British Pound could be classified as being the same asset but it is left open whether both currencies can actually be classified as being the same asset by looking at fundamental variables. To the best of our knowledge, there is no contribution in the literature, apart from Hakkio and Rush (1989) and Baffes (1994), who only discuss this point verbally, as well as Ferré and Hall (2002), who look at only one common stochastic trend, that theoretically investigates currencies for being the same asset as indicated by strong-comovements of exchange rates.

Departing from the assumption of obtaining stationarity by differencing one time, i.e. the levels are integrated of order 1, Dwyer and Wallace (1992) point out that if two exchange rates in terms of the same base currency are cointegrated the cross rate is required to be stationary. Baillie and Bollerslev (1994) argue by using the methodology of Granger and Joyeux (1980) that exchange rates (in a multivariate framework) are fractionally cointegrated, i.e. the integration parameter of the residuals is a fraction. In a bivariate cointegration analysis the cross rate is the residual and hence must be either stationary or mean-reverting fractionally integrated. Cheung (1993) demonstrates for five exchange rates in terms of U.S. dollars by using the Geweke and Porter-Hudak (1983) (GPH) estimator that the observed exchange rates are fractionally integrated, i.e. following a long-memory process. In contrast, Barkouklas et al. (2004) do not find long memory in exchange rates. The most recent contribution in this respect, Jin et al. (2006), explores 19 exchange rates denominated in U.S. dollars and finds fractionally integration for most of them.[4] Granger (1986) introduces the concept of fractional cointegration that is applied by Cheung and Lai (1993) to the purchasing power parity. In a triangular context, fractional integration of one exchange rate would be evidence of fractional cointegration if the other two involved exchange rates are integrated of the same order. As suggested by Baillie and Bollerslev (1994), it is possible that more exchange rates are bivariately fractionally cointegrated and behave strongly linked over the long-run indicating that common fundamentals are relevant in both exchange rates.

The aim of this paper is to close three gaps in the literature. Firstly, a theoretical investigation of reasons why exchange rates strongly comove apart from Hakkio and Rush (1989), Baffes (1994) and Ferré and Hall (2002) is missing. Secondly, no contribution has explicitly taken economically and empirically the empirical results of Sephton and Larsen (1991) into account and ascertain why the finding of cointegration of exchange rates heavily depends on the selected period. Thirdly, the same can be said with respect to efforts to determine periods in which exchange rates are cointegrated.

[3] If the hypothesis of no cointegration is rejected and the cointegration vector is not consistent with the no-arbitrage condition, the market is seen to be (informationally) inefficient because the currencies expressed in the same currency are different asset but their exchange rates comove over the long-run.

[4] Cheung (1993) uses weekly data while Jin et al. (2006) focus on monthly data.

Therefore, we scrutinize theoretically and empirically periods in which two currencies can be classified as being the "same" asset. The empirical results yield that periods of strong comovements between the U.S. dollar and the Pound sterling based upon the Euro prevail during the 1990s and periods of comovements between Euro and Pound sterling denominated in U.S. dollar prevail since the introduction of the Euro. Furthermore, no long-run relationships can be discovered.
The remainder is organised as follows. After the introduction, section 2 provides theoretical considerations on when currencies can be classified as being the same asset. Firstly, a general discussion is provided by employing the asset approach of exchange rates. In section 3 the econometric methodology is outlined. In the following section 4 the empirical results are presented. An economic interpretation of the empirical results is carried out in section 5. Section 6 finally concludes.

2 Econonomic framework

2.1 The asset approach on exchange rates

The question to answer is under which conditions different currencies can be seen as the same asset. From Hakkio and Rush (1989) and Baffes (1994) follow that two currencies are not different assets if their exchange rates vis-à-vis the same denomination currency are cointegrated. Two currencies become the same asset if a set of fundamentals is common to both exchange rates.[5]
In this section the conditions are derived under which a cointegration relationship between two exchange rates denominated in the same currency is established if all three relevant fundamentals' processes are allowed to be non-stationary. In order to assess the impact of fundamentals on exchange rates, a very general framework is applied. Since the underlying question is concerned with the inspection of exchange rates as same assets, the "asset approach" by Mussa (1976) and Frenkel and Mussa (1980) is used. Thus, we maintain the consistency to investigate currencies as assets.
Exchange rates should be seen as relative prices of national monies instead of relative prices of national outputs.[6] For this reason, the determination of the exchange rates does not take place under the condition of equilibrium in the markets for flows of funds. Since relative prices of domestic monies are concerned, the conditions for equilibrium in the markets for stocks of assets are relevant. The monetary approach to exchange rates by Frenkel (1976) is an example of that kind of models. Here, the equilibrium results from the demand for and the supply of domestic money which means the stocks of both counterparts are relevant. The exchange rate determination equation is given by

$$s_t = (m_t - m_t^*) - \alpha_1(y_t - y_t^*) + \alpha_2(i_t - i_t^*) \tag{1}$$

with m as money supply, y as real income and i as the interest rate, whereas $*$ denotes foreign variables. α_1 and α_2 denote the (semi-) elasticities.[7]

[5] Further, Baffes (1994) writes that this can be the case if the two countries whose currencies are expressed in the same currency fix their bilateral exchange rate or the economic structure and policy completely coincide. C.f. p. 275.

[6] Frenkel (1976), Dornbusch (1976), Kouri (1976), and Mussa (1976) treat the foreign exchange market under a similar perspective in an issue of the Scandinavian Journal of Economics of 1976. While Dornbusch (1976) highlights the importance of the money market for short-run developments Kouri (1976) looks at the process of asset accumulation. Frenkel (1976) stresses the monetary approach. Mussa (1976) provides the seminal arguments for the asset view of exchange rates.

[7] The higher the domestic (foreign) money supply the higher the domestic (foreign) price level is, thus resulting in a depreciation of the domestic (foreign) currency. Another relevant factor on the money market is income. An

In the face of disequilibrium on the money market, subsequent flows of funds are generated to restore the equilibrium but these flows are not important for determining the exchange rate (Mussa, 1976, pp. 233-234). As in the asset approach for stocks changes in exchange rates are largely unpredictable. Thus, occurrences of new information have an impact on expectations regarding the underlying fundamentals and hence are the driving forces in the exchange rate determination process. From this point of view, rather than the magnitude of flows it is the willingness to hold stocks at each market condition that is important for the pricing process (Mussa, 1976, p. 235).[8]

The monetary approach is special case of the asset approach and it can therefore be formulated a generalized model that takes account for expectations and in which the question of the variables' choice is not pre-specified. It is assumed that the exchange rate at time t (s_t) depends on the initial environmental conditions contained in the vector x_t (all expressions in logarithms) and on the expected change of the current exchange rate to the next period, i.e. the difference between the current and the expected exchange rate of the consecutive period based on the information set Φ_t available at the current period.

$$s_t = x_t + b \cdot [E(s_{t+1}|\Phi_t) - s_t] \tag{2}$$

The initial environmental conditions describe the fundamentals in period t and hence also the supply and demand coming from the resulting market conditions. The second part of the exchange rate equation arise from the assumption that the market players form expectations regarding the exchange rate they expect in the next period ($E(s_{t+1}|\Phi_t)$). To be complete, the parameter b weights the impact of the expected change (Frenkel and Mussa, 1985). The greater the value of b, the more weight lays on the expected change, i.e. the more sensitive the exchange rate will be to expected changes.

At this point, it is assumed that the expectations of the market players are rational.[9] Thus, the market players know the validity of equation (2) in each subsequent period. To be more precise, the expected exchange rate in $t+1$ depends on the basic conditions in time $t+1$ (if already known or expected) and the expected change in the exchange rate from period $t+1$ to $t+2$ conditioned on the information set in t, whereas the weight b is held constant. For all consecutive periods the same consideration can be applied. Through forward iteration and resolving for s_t the exchange rate formation equation can be derived. Assuming the validity of the transversality condition[10], i.e. ruling out rational bubbles, equation (3) is obtained.

$$s_t = \frac{1}{1+b} \sum_{k=0}^{\infty} \left(\frac{b}{1+b} \right)^k \cdot E(x_{t+k}|\Phi_t) \tag{3}$$

From equation (3) it results that the exchange rate in period t is the discounted sum of fundamentals' expected values conditioned on information set Φ_t, i.e. on the information available at time t, the

increase in income raises the money demand. Holding all other variables constant, the excess demand can only be equilibrated by increasing real money supply which is achieved if the price level decreases. A similar argumentation can be applied to the interest rates. Raising interest rates reduces money demand resulting in an excess supply. Money market equilibrium can only be restored c.p. if real money supply decreases which happens when the price level rises (here, there is no economic causality, only conditions for market equilibrium).

[8] At this, the market volume is not important for the extent the prices change. It is more important the desirability to hold stocks at each relevant market price, i.e. exchange rate. For this reason, Mussa (1976) argues about the importance of sentiments' structure. The trade volume only depends on the degree of accordance in sentiments.

[9] Expectations are rational in the sense that the market players know the relevant equilibrium pricing model and they also are able to use all available information correctly.

[10] C.f. Blanchard and Fischer (1989), p. 43 for transversality condition and Chapter 5 for bubbles.

current period (Frenkel and Mussa, 1985). The general solution of the asset approach is closely linked with the efficient market theory. A market is seen to be efficient if all available and for the pricing process relevant information is contained in the current exchange rate (Fama, 1970; LeRoy, 1989). If the market forms rational expectations regarding future exchange rates (or based upon expected fundamentals), i.e. the validity of equation (3), all information is already included in the exchange rate and no excess profits can be earned by exploiting the current information set (Jensen, 1978). Changes in expected fundamentals are already included in the current exchange rates.[11] Hence, only new information about future fundamentals can alter the exchange rate (Mussa, 1976). The extent to which the exchange rate changes in the face of new information about future fundamentals depends on the discount factor in equation (3).

2.2 Currencies as the same asset

The above generalized outlined theory of exchange rate determination shall be used to evaluate the conditions under which exchange rates move closely together in the long-run. For this reason, we treat a three country-three currency case in which all exchange rate are assumed to be flexible. The corresponding exchange rate equation is provided in equation (4) for country 2 and in equation (5) for country 3, whereas the currencies are denominated in terms of currency 1.[12]

$$s_t^{12} = \frac{1}{1+a} \sum_{k=0}^{\infty} \left(\frac{a}{1+a} \right)^k \cdot E(x_{t+k}^{12} | \Phi_t^{12}) \tag{4}$$

and

$$s_t^{13} = \frac{1}{1+b} \sum_{k=0}^{\infty} \left(\frac{b}{1+b} \right)^k \cdot E(x_{t+k}^{13} | \Phi_t^{13}) \tag{5}$$

Since the asset approach treats the exchange rate as the relative price of domestic assets, two destinations have an impact on the exchange rate. Thus, the processes of fundamentals in equations (4) and (5) must be split up in each case into two processes covering the fundamentals of the first and the second country x_t^i and x_t^j

$$x^{1j} = x_t^1 - x^j \tag{6}$$

with $j = [2,3]$. By using equation (6) and inserting in equation (4) it results equation (7)

$$\begin{aligned} s_t^{1j} &= \frac{1}{1+a} \sum_{k=0}^{\infty} \left(\frac{a}{1+a} \right)^k \cdot \left[E(x_{t+k}^1 | \Phi_t^1) - E(x_{t+k}^j | \Phi_t^j) \right] \\ &= \underbrace{\frac{1}{1+a} \sum_{k=0}^{\infty} \left(\frac{a}{1+a} \right)^k \cdot E(x_{t+k}^1 | \Phi_t^1)}_{F_t^1} - \underbrace{\frac{1}{1+a} \sum_{k=0}^{\infty} \left(\frac{a}{1+a} \right)^k \cdot E(x_{t+k}^j | \Phi_t^j)}_{F_t^j} . \end{aligned} \tag{7}$$

For the sake of simplicity, the two processes of discounted expected fundamentals can be abbreviated by F_t^1 for the first country and F_t^j for country j. In general, the time series properties of the fundamental series are important for the evaluation of the time series properties of the exchange

[11] The expected change can be expressed as the discounted difference between the expected exchange rate in the next period based upon the information in t and the initial environmental fundamentals, x_t. C.f. Frenkel and Mussa (1985), p. 727.

[12] The sensitivity parameter for s_t^{12} is a and for s_t^{13} b.

rate series. As Engel and West (2004, 2005) show, one case claims more attention. If the discount factor in equation (2) approaches zero, i.e. only a small weight lies on the expected change, and if the fundamentals are first-difference stationary, then the exchange rate exhibits near-random walk behaviour.[13] This case is also interesting and should be borne in mind.

More attention should be directed to equations (8) and (9).

$$s_t^{12} = F_t^1 - F_t^2 \tag{8}$$

and

$$s_t^{13} = F_t^1 - F_t^3 \tag{9}$$

In the first case, the exchange rate is simply the difference between the process of fundamentals F_t^1 and F_t^2 and in the second case respectively.

In a triangular framework the market is efficient if no cross market arbitrage opportunities exist. Frenkel and Levich (1975) and Levich (1985) speak about triangular arbitrage in this context meaning that if the exchange rates are expressed in logs the difference between two exchange rates must be equal to the third one (equation (10)).[14]

$$s_t^{12} - s_t^{13} = s_t^{32}. \tag{10}$$

The two currencies under observation (currencies 2 and 3) are the same asset if their exchange rates denominated in the same currency (currency 1) are cointegrated with cointegration vector $(1, -1)$. In addition to equations (8) and (9), equation (10) is valid. Thus, equations (8) and (9) can be set in equation (10) and it results

$$s_t^{12} - s_t^{13} = (F_t^1 - F_t^2) - (F_t^1 - F_t^3). \tag{11}$$

In order to be cointegrated, it is required that a linear combination of s_t^{12} and s_t^{13} is stationary. The left-hand side of equation (11) is the linear combination if the no-arbitrage condition holds. Hence, the right-hand side must also be stationary. A short reformulation yields the result that the process of fundamentals with respect to country 1 cancels out. It is straightforward to argue that if the exchange rates are cointegrated and the no-arbitrage condition holds their difference is equal to the cross rate and therefore equal to the difference between the fundamentals of country 3 and 2.

$$s_t^{12} - s_t^{13} = (F_t^3 - F_t^2) = s_t^{32} \tag{12}$$

Since cointegration requires the left- and right-handside of equation (12) to be stationary, it can be seen that stationarity only arises if either both processes of fundamentals are stationary or if the process of fundamentals of country 2 and 3 are cointegrated.

This is consistent with the intuition as seen by the asset approach, the domestic currencies are expressed in the same currency, and hence, *their* fundamentals (country 2 and 3) are also converted into the same currency (of country 1). This result arises independently from the choice of the denomination currency. When currencies 2 and 3 are expressed in a fourth or fifth currency, these

[13] In Engle and West (2004, 2005) equation (2) is expressed differently. Hence, their discount factor approaches one.

[14] If transaction costs are present in the market the equality is to be adjusted by a constant terms that drives a wedge between the difference and the cross rate.

exchange rates are also cointegrated. Both currencies are the same asset if their fundamentals are linked in the long-run. In other words, neglecting short-run disturbances the process of the fundamentals has common stochastic trends.[15]

3 Econometric methodology

3.1 Tests for stationarity

It is assumed that the time series y_t can be expressed with the help of the following $ARIMA$ representation:

$$(1-L)^d A(L) y_t = C(L)\epsilon_t{}^{16} \tag{13}$$

Since the white noise error term ϵ_t is stationary, d is the number with which the left-hand side must be differenced to obtain stationarity. The time series y_t is integrated of order d, i.e. $y_t \sim I(d)$. At this point, the parameter d shall only be a non-negative integer. For $d = 1$ the time series is non-stationary meaning that a disturbance rests indefinitely in the series. In the case of $d = 0$ the time series is stationary and an innovation dissipates over time, so that the series wander around a specific mean value.[17] The autocorrelation function of the process decays exponentially within a very short period. Stationarity must be correctly denoted by covariance-stationarity. This means that not only the mean is independent of time but the covariances and the variance are, too. Hence, for $d = 0$ the time series is stationary and mean reverting.

The parameter d does not need to be a non-negative integer. As Granger (1980), Granger and Joyeux (1980) and Hosking (1981) demonstrate, d can also be a fraction, i.e. any real number. Again by using equation (13), the term $(1-L)^d$ is defined more generally. In this context, one speaks about fractional integration.[18] Hence, if the parameter d is a non-integer, the time series under observation is fractionally integrated.[19] For this reason, the integrated autoregressive moving average ($ARIMA$) model in equation (13) becomes a fractionally integrated autoregressive moving average ($ARFIMA$) model. In that case, the autocorrelation function decays hyperbolically. A disturbance has a long-lasting effect and only dissipates over a long-run horizon whereas it dies out quickly for pure $I(0)$ processes and never for non-stationary processes. For this reason, the case of fractional integration is also referred to as long-memory. Following Hosking (1981), the series is stationary if $0 \leq d < \frac{1}{2}$ (pp. 169-171). If the integration parameter lies within the interval $\frac{1}{2} \leq d < 1$ the time series is not stationary any longer but mean-reverting. A disturbance still dies out in the long run; only the covariances depends on time (Baillie, 1996, pp. 21/22).[20]

Usual tests that exclude the possibility of fractional integration can thus only distinguish between $I(1)$ (non-stationarity) and $I(0)$ (stationarity). The classical tests for testing stationarity depart

[15] If more than one process of fundamentals is used - as usual, one can argue that the fundamentals have to be panel-cointegrated for cointegration between the exchange rates.

[16] ϵ_t is a white noise error process. L represents the lag operator and d is the parameter of integration. $A(L)$ and $C(L)$ are the polynomials in the lag operators of y_t and ϵ_t respectively with roots outside the unit circle.

[17] In eq. (13) the mean is not different from zero.

[18] As given in Granger and Joyeux (1980) and Hosking (1981), the function $(1-L)^d$ can be expressed in terms of a gamma function Γ: $(1-L)^d = \sum_{k=0}^{\infty} \frac{\Gamma(k-d)L^k}{\Gamma(-d)\Gamma(k+1)}$.

[19] d can also have negative values but this case shall be excluded here.

[20] The definitions for long memory models vary depending on the field in which they are used. A broad literature review is provided by Baillie (1996) including a deeper statistical discussion.

from the Dickey-Fuller (DF-) test and start with a simple regression of the change in y_t on its one period lagged levels (Fuller, 1976; Dickey and Fuller, 1979).[21] In order to account for serial correlations, the original DF test is extended by allowing for lagged independent variables of a higher order (augmented Dickey-Fuller (ADF-) test). A problem with ADF tests is that the choice of the lag length must be done very carefully because serial correlations can distort the test statistics as well as heteroskedasticity can do. Phillips (1987), Perron (1988), and Phillips and Perron (1988) develop a semi-parametric test that can deal with serial correlations and heteroskedasticity. The critical values of this test are equal to those of the ADF test.[22] In the PP-test, the lagged variables on the right-hand side are omitted because the serial dependencies are caught by using the robust Newey-West estimator. The ADF and the PP test have low power when a moving average is important. For this reason Elliot et al. (1996) revised the ADF test insofar as they use a GLS instead of an OLS estimator. These tests yield more efficient results than the ADF and the PP-test. A unit root test with an alternative formulation of the null hypothesis is provided by Kwiatkows et al. (1992). Here, it is tested for stationarity against non-stationarity under the alternative.

An approach for the estimation of fractional integrated processes is provided by Geweke and Porter-Hudak (1983) (GPH hereafter). This approach can estimate the parameter d sufficiently for values $d < 1/2$ and is the most popular technique to detect long-memory behaviour of time series. GPH depart from the periodogram (finite sample estimate of the spectral density) that is closely linked to the autocorrelation function of the process.[23] The idea of the GPH estimator is to separate the low-frequency behaviour of a time series from higher frequency components and to investigate the series at harmonic frequencies. In that case, the integration parameter can be estimated by OLS with the regression of the log periodogram at a specific ordinate on a constant and on the harmonic ordinates. The parameter d is then the slope coefficient of the harmonic ordinates (Geweke and Hudak-Porter, 1983, p. 225). When d is zero, none of the periodogram can be explained by the frequency at the specific level of the periodogram.[24] A major drawback of the log periodogram regression in the fashion of GPH is that the estimate $\hat{d}$ is inconsistent when the true value of d is greater than one, i.e. in the unit root case (Kim and Phillips, 2006).[25] Phillips (2007) revises the GPH estimator by using an exact log periodogram regression and additionally develops an asymptotic theory for unit root cases.[26]

3.2 Tests for Cointegration

As discussed in the previous section, the cointegration methodology is applied for the investigation when currencies can be seen as same assets. We have already pointed out that two non-stationary time series are cointegrated if a linear combination of them exists that is stationary (Engle and

[21] To be more precise, these are no tests for stationarity, rather for a unit root, i.e. the null hypothesis is non-stationarity.

[22] Exact critical values and p-values for the ADF test are provided by MacKinnon (1996) who uses a response surface regression for their computation.

[23] See Granger and Joyeux (1980) and Hosking (1981) for a more detailed discussion.

[24] The least squares regression equation is $\ln(I(\omega_j)) = c - d \cdot \ln(4\sin^2(\omega_j/2)) + error_j$ with $j = 1, ..., n$ and $\omega_j = 2\pi j/T$ where $j = 1, ..., T-1$. $I(\omega)$ is the periodogram of the series under observation at frequency ω_j. The spectrum of the time series Y $(I(\omega))$ is defined as $\left|1 - e^{-i\omega}\right|^{2d} I_u(\omega)$ where $I_u(\omega)$ is the spectrum of $u_t = (1-L)^d Y_t$. See Granger and Joyeux (1980), Hosking (1981), Geweke and Porter-Hudak (1983), Agiakloglou et al. (1993), and Baillie (1996).

[25] When the true process is non-stationary, Hurvich and Ray (1995) quantify the bias of the GPH estimator.

[26] See Phillips (2007) for a more detailed discussion, derivations and proofs. One substantial change affects the number of frequency ordinates. GPH propose $m = O(T^{1/2})$. Phillips (2007) provides a more rigorous $m = O(T^{1/2-1/p})$.

Granger, 1987).[27] Furthermore, some time series are contained in the vector X_t. The elements of X_t are cointegrated of order d, b ($X_t \sim CI(d,b)$), if they are all integrated of the same order and a vector exists such that

$$Z_t = \beta' X_t \sim I(d-b) \tag{14}$$

with $b > 0$.[28] In this multivariate case, the vector β_i is the cointegration vector and contains the cointegration parameters (Engle and Granger, 1987, pp. 252-255).

The original way to test for cointegration bases upon equation (??). Engle and Granger (1987) propose to run a linear regression and to test the resulting error series for stationarity. A second technique developed by Johansen (1988, 1991) makes use of equation (14) as well as Granger's representation theorem that states that a cointegration relationship is equivalent with an error correction representation. Johansen departs from a vector autoregressive model and formulates it as a first difference VAR with an error correction term

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-1} + \epsilon_t, \tag{15}$$

whereas $\Gamma_i = -I + \Pi_1 + ... + \Pi_i$ with $i = 1, ..., k-1$ and $\Pi = -(I - \Pi_1)$ (Johansen, 1988, 1991; Johansen and Juselius, 1990). The matrices Γ_i include information about short run adjustment coefficients. The expression ΠX_{t-1} denotes the error correction term and contains the cointegration relationship. A test for cointegration is directed to test for the rank of ΠX_{t-1}. The rank of a matrix is the number of independent columns and is here automatically equivalent with the number of cointegration vectors. Thus, the rank of ΠX_{t-1} shows the cointegration rank and if Π is not of full rank it can be decomposed into two matrices α containing the adjustment coefficient to long-run equilibrium and β, the cointegration vector, with dimension $p \times r$ whereas $r \leq p$ such that $\Pi = \alpha\beta'$.

Granger (1986) introduces the concept of fractional cointegration which is implicitly contained in the generalisation of Engle and Granger (1987).[29] Cheung and Lai (1993) provide a practical application of the concept of fractional cointegration on exchange rates and the validity of purchasing power parity. Time series are fractionally cointegrated if the equilibrium error's integration parameter is less than 1. Then, a shock to the system dies out over the long-run, so that the equilibrium relationship between the time series is retained (p. 106). Furthermore, Cheung and Lai (1993) argue that the least squares estimate is consistent in the case of fractional cointegration and that only the rate of convergence differs.[30] As an estimator they applied the GPH method to the residual series.

For this reason, a test for fractional cointegration is a test for $I(1)$ versus $I(d)$ with $d < 1$. From this point of view, the Phillips (2007) estimator can also be used to test for (fractional) cointegration of exchange rates within a triangular framework. If the estimate of d is significantly different from zero and smaller than 1, two exchange rates are fractionally cointegrated inasmuch they are integrated of the same order.

[27] For a closer explanation see Granger (1981) and Granger and Weiss (1983).

[28] d denotes the number of independent cointegration vectors.

[29] Granger (1986) refers to Granger and Joyeux (1980) and argues that fractional differencing can be applied and the general representation still holds (p. 222). Engle and Granger (1987) formulate the general defintions but concentrate only on the $I(0)$ and $I(1)$ cases (pp. 252/253).

[30] Cheung and Lai (1993) show the consistence of the least squares estimate analytically (p.106).

Granger and Ding (1996) and Granger and Hyung (2004) depart partly from the notion of long-memory behaviour of financial time series. In both contributions, it is pointed out that fractional integration can arise from occasional structural breaks. Instead of using a fractional integration framework more emphasis shall be laid on the modelling of structural changes. We take account of this critique by using rolling regressions. The PP and the KPSS tests as well as the modified log periodogram regression are performed with a fixed sample size and are rolled over the whole period of observation.

4 Empirical results

4.1 Data

For the purpose of detecting periods in which exchange rates are the same assets, the weekly end of day data for the four most traded currencies are taken into account.[31] The used currencies are the Euro, the US dollar, the Japanese yen and the British pound sterling.[32] Between these currencies each exchange rate is examined. Before the introduction of the Euro in January 1999, the Deutsche Mark (DEM) rates are used and converted with the official DEM/Euro conversion rate into Euro. While doing so, the DEM is not seen as the economic predecessor of the Euro. The reason why the DEM is used traces back to the fact that the DEM was the second most traded currency before the introduction of the Euro. From this point of view, the Euro replaced the DEM in the quantitative importance. For sake of simplicity, it is referred to the Euro rates and the Euro area when the DEM and Germany are investigated in periods up to 1998. All exchange rates are expressed in natural logs and taken from Datastream. The period of observation starts on 5 January 1994 after the turbulences caused by the EMS crises diminished and ends on 29 June 2007 before the sub-prime crises broke out.[33]

4.2 Unit root tests

Since the aim of this paper is to detect periods in which exchange rate pairs are cointegrated, i.e. currencies are the same asset, it must be checked for the premises of cointegration. For this reason, all exchange rates in logs are tested for stationarity. This is done for the whole period of observation but the sample is also subdivided into the period before and after the introduction of the Euro. The results for the stationarity tests are reported in tables 2, 3 and 4. According to the above outlined unit root tests, the PP, KPSS and the DF-GLS tests are applied to the levels and first differences. If the PP and DF-GLS tests can reject the null hypothesis for first differences but not for levels, it is concluded that the exchange rates under investigation are first difference stationary. For the KPSS test the reverse argumentation applies. A non-rejection of the null hypothesis for first differences is evidence for first difference stationarity when it can be rejected for levels. As can be seen in table 2, the stylized facts of Meese and Singleton (1982) and Baillie and Bollerslev (1989b), that exchange rates are first difference stationary, can be confirmed. For the exchange rates EUR/USD and GBP/USD the DF-GLS test cannot reject the null hypothesis of non-stationarity in first differences. Since the KPSS test cannot reject the null hypothesis of stationarity based

[31] The exchange rates are averages of the bid and ask rates.

[32] Together these exchange rates make up a share of approximately 160 % of the average daily turnover. The share is greater than 100 % because it is measured the turnover in which the currency is involved.

[33] In this period significant foreign exchange market interventions are only expected when the JPY is concerned.

upon robust grounds, these exchange rates are treated as first difference stationary. In the second sub-period the KPSS test cannot reject the null hypothesis for levels while the DF-GLS test can at 5% significance level for the JPY/USD rate. Here, evidence that the JPY/USD is stationary seems weak.

Although the KPSS-test is able to distinguish between stationarity and cases in which $-1/2 < d < 1/2$ (Lee and Schmidt, 1996) it is known that Dickey-Fuller type unit root tests are not robust when the true data-generating process is fractionally integrated (Diebold and Rudebusch, 1991). For this reason, more attention shall be directed to the integration parameter. At this, we apply the estimator developed by Phillips (2007) that is a variant of the GPH estimator. Jin et al. (2006) is the most recent contribution that investigates exchange rates for long memory and they use the so called OLS wavelet estimator. They compare the estimates of their estimator with those of the GPH estimator and conclude that the results are completely disparate (p.132). While the wavelet estimator gives evidence in favour of long-memory, the GPH estimates contrast these results. For this point of view, they assert the theoretical investigation by Jensen (1999) who developed the OLS wavelet estimator. Jensen (1999) shows that the mean squared error of his estimator is significantly smaller than that of the GPH estimator. In addition, the GPH based tests are biased positively away from zero (Agiakloglou et al., 1993) whereas the quantity of the bias is negligible (Liebermann, 2001).[34] In order to test for fractional integration, the modified log periodogram regression by Phillips (2007) is applied because it is able to determine the order of integration over a broad interval. Although Phillips (2007) also mentions the lower consistency of his estimator we decide to apply his test anyway because of the attractiveness of estimating d in unit root cases. We are aware of the critiques and thus use the modified log periodogram regression in addition to the basic unit root tests.

In Cheung (1993), Barkoulas et al. (2004), and Jin et al. (2006) the (log) changes of the exchange rates are inspected for long-memory behaviour.[35] Their studies are focussed on cases in which $-0.5 < d < 0.5$.[36] Since our aim is to identify periods in which the level series are stationary and/or mean-reverting, we analyse the (log) levels of exchange rates. The modified GPH test by Phillips (2007) is able to test for the case of $I(1)$. The application of the exact log periodogram regression (ELP) is given in table 5. The hypothesis that d is equal to 1 can broadly be confirmed for all periods. Only in the case of GBP/USD, is there evidence for fractional integration. An increase of the power of the test from the usual 0.5 to 0.6 changes the result and provides evidence in favour of $I(1)$ in all cases.[37]

Thus, it can be concluded that all exchange rates can be used to perform a cointegration analysis. They are all integrated of the same order. The above outlined economic argumentation departs from the assumption that the no-arbitrage condition holds. Hence, it must first be carefully tested whether this is true in reality. The no-arbitrage condition in eq. (10) can be slightly rearranged in

[34] Agiakloglou et al. (1993) show for moderate sample sizes (100 observations) that the estimator of d is substantially biased when the autoregressive parameters have large positive values. In general, the hypothesis of pure stationarity is too frequently rejected. Furthermore, Agiakloglou et al. (1993) argue that the bias is only a serious problem when the autoregressive parameters are large and positive (pp. 236/237). The same is true for large positive values of moving-average parameters (pp. 240-242). Lieberman (2001) quantifies the bias and concludes that the GPH estimates are sufficiently reliable.

[35] In this way, exchange rates returns are investigated. Granger and Hyung (2004) scrutinize also the return series.

[36] In Barkoulas et al. (2004) only the case of $0 \leq d < 0.5$ is treated.

[37] This result is not printed here but available from the author.

bringing the right-hand side to the left-hand side.

$$s_t^{12} - s_t^{13} - s_t^{32} = \mu \tag{16}$$

Equation (16) states that no excess profit can be earned after cross converting. When transaction costs as defined above are relevant (μ), the left hand-side of equation (16) is equal to a constant term. As Baffes (1994) argues, the no-arbitrage condition is fulfilled when the three exchange rates are cointegrated. This is tested by applying the Johansen (1988, 1991) approach to each triplet.[38] In all cases, the validity of the no-arbitrage condition can be shown (see table 6).[39]

The validity of the no-arbitrage condition makes possible to conclude that two exchange rates denominated in the same currency are cointegrated when they are non-stationary while the cross rate is stationary. A test for cointegration is therefore equivalent to a test for stationarity of the cross rate.[40] By reviewing the results of the unit roots tests, it can be concluded that no cointegration of exchange rates is expected over the whole period of observation because all exchange rates are integrated of the same order.

4.3 Testing for stationarity applying a rolling regression framework

Since there is evidence that structural changes seem to play a significant role in answering the question whether currencies can be seen as the same asset in reality based upon the observations of Sephton and Larsen (1991) and Kühl (2007), we return to the basic unit root tests and the ELP. Their analyses have shown that cointegration seems to be present in shorter periods. In the following, we try to detect those periods in which exchange rates are cointegrated. For this reason, the KPSS and the PP test as well as the ELP estimator are applied within a rolling regression framework. The estimation period is held fixed at a size equivalent to two years (96 observations). This period is rolled over the whole period of observation and at each time the test statistics are reported. In contrast to a rolling window regression framework, a regime swichting approach could also be applied. Since the cointegration analysis is a long-run concept and regime switching approaches can pervert the idea of cointegration when the regimes last to short we decide to fix exogenously the window. We are aware of the critique of Hakkio and Rush (1991) who argue that it is the time span that should be extended to account for a long-run equilibrium concept rather than the frequency and take it into account to the extent that we rely on weekly data. The reason for the size of the rolling window is that the length of two years is an appropriate point of departure when looking at long-run relationships on the foreign exchange market.

The visual results are presented in figures 1, 2 and 3. For the PP-test the p-values based upon the exact critical values of MacKinnon (1996) are used to evaluate the significance of the test statistic. If the calculated p-value is less than the corresponding significance level, the null hypothesis of non-stationarity is rejected. A similar argumentation can also be applied to the ELP estimator. Here, there is evidence in favour of stationarity if the p-value exceeds the corresponding significance

[38] Following the Johansen definition of cointegration, it is not required that all variables are integrated of the same order as long as they build an error correction term.

[39] There is one caveat: the used exchange rates are averages of bid and ask rates. Economically, bid rates should be used because they are relevant by converting one currency in another. Unfortunately, in all cases in which the DEM/JPY rate is involved only average rates are available. Due to the lack of data, the available average data are used. For all other exchange rates the same tests are applied by using bid rates and the results do not differ. Thus, it is expected that the same is true for the DEM/JPY, although the doubts cannot be ruled out completely.

[40] Needless to say, this approach cannot be applied in a multivariate framework with more than two exchange rates denominated in the same currency.

level. The p-value corresponds with the null hypothesis that d is zero. The higher the p-value is, the more robust the result is in favour of stationarity. By using the KPSS test, the critical values are reported and compared with the critical values of the corresponding significance level.
Based upon the tests, three different periods can be identified. The first starts in the mid 1990s. Here, the GBP/USD exchange rate is stationary while for example the EUR/USD and EUR/GBP are non-stationary.[41] After the introduction of the Euro, the PP-test again shows stationarity of EUR/USD, GBP/USD and EUR/GBP rates simultaneously, which is also confirmed by the KPSS test. From this point of view, the currencies involved are expected to be the same assets from round about end 1999 or beginning of 2000 till mid or end of 2002, i.e. the exchange rates denominated in a fourth currency and not belonging to the triplet should be cointegrated. The third period starts in mid-2004. Here, the EUR/GBP exchange rate seems to be stationary again while the other exchange rates are non-stationary.[42]
The outlined results are remarkably robust because they do not change by applying a sensitivity analysis. The fixed sample within the rolling regression framework is extended to three years (144 observations). Although the length of consecutive verification of our hypothesis becomes smaller, the same quantitative results with respect to the sub-period arise. The same is true for the ELP estimator when the power of the estimator is increased to 0.6.
Finally, it can be concluded that exchange rate pairs comove during subsequent periods and the involved currencies can be seen as the same asset.

5 Economic analysis

5.1 Classifications of the periods

Since the most interesting behaviour occurs in the triplet EUR/USD, GBP/USD and EUR/GBP, these weekly exchange rates are presented in figure 4, whereas the periods of stationarity are shaded.[43] The periods in which the above mentioned rolling regressions can find stationarity for at least one exchange rate are very similar across the techniques. All three tests indicate approximately the same point in time when the period of stationarity starts; only the duration differs. For the further purpose of the investigation, the KPSS test results are predominantly used for the categorisation of periods because this test is fairly robust. As already pointed out, three different regimes are classified. The first one starts approximately at the beginning of October 1996 and ends around mid May of 1999. In this period the GBP/USD exchange rate is stationary and hence the EUR/USD and EUR/GBP rates are cointegrated. Although the visual inspection suggests evidence that a comovement between the EUR/USD and EUR/GBP rates is already present before the classified period, none of the tests can confirm that result statistically. Hence, this comovement is not that strong as indicated by cointegration.[44] This period is referred to as period 1. By looking at the classification of the sub-sample obtained by the KPSS-test, the EUR/GBP exchange rate is stationary in the time interval mid October 1999 till mid April 2002. The period of stationarity for

[41] The PP test exhibits stationarity for these exchange rates, too. But the two other tests do not. This allows us to conclude to treat only one exchange rate as stationary.
[42] In the case of the EUR/GBP, Kühl (2007) investigates the period after the introduction in more detail and finds by using the Johansen approach cointegration between EUR/USD and GBP/USD in sub-periods. His investigation is directed to market efficiency considerations.
[43] The exchange rates are normalised to the first observation, i.e. first week in January 1994.
[44] The PP test and the ELP test indicate cointegration slightly earlier than the KPSS test does. Qualitatively, the broader classification has no consequences.

the EUR/USD exchange rate runs from the end of June 2000 till the end of June 2002. Finally, the GBP/USD rate is stationary from mid August 2000 till the beginning of October 2002. The overlapping period in which all three exchange rates are stationary is labelled period 2 and runs from mid August 2000 till mid April 2002. We could investigate each exchange rate separately but in this case much of the dynamic between the variables would be lost. Thus, an exploration of joint effects would be appropriate. But the number of observations in this period is less than our exogenously given 96 observations equivalent to two years in the rolling regression framework. This is the reason why we decide to neglect this period in the following economic analysis as it does not match our definition of the long-run. The third period starts at the end of June 2004 and runs till mid June 2006. In this period the EUR/GBP exchange rate is stationary and because the EUR/USD and GBP/USD rates are non-stationary, they are cointegrated. The visual inspection gives evidence that the period of stationarity for the EUR/GBP exchange rate lasts longer than indicated by the KPSS-test. It seems to match with the results of the PP-test and the log-periodogram regression. Nevertheless, we maintain the classification achieved by the KPSS-test.

5.2 Analyzing the cointegration relationships

To get a deeper insight into the linkage between the exchange rates, we employ for both periods the Johansen cointegration analysis to those exchange rates that are considered to be cointegrated as indicated by the rolling regressions. Since the sample size is very small and the trace statistic tends to overreject the null hypothesis in small samples as discussed by Cheung and Lai (1993), we use the small sample correction as proposed by Johansen (2002). In both cases, the null hypothesis of no cointegration can clearly be rejected as given in tables 8 and 9. In the first sub-period the null hypothesis is rejected with a p-value of 0.003. But for the pairings EUR/USD and EUR/GBP the null hypothesis of rank one can also be weakly rejected. This would mean that all exchange rates are stationary. Since our previous analysis rules out this case, we assume that there is just one stationary linear combination. The rejection of zero rank in the third sub-period takes place with a p-value of 0.039.

The unrestricted estimation of the vector error correction model shows that the EUR/USD exchange rate is weakly exogenous because the hypothesis that the adjustment coefficient is not different from zero cannot be rejected as indicated by the t-statistic of -1.368.[45] Hence, the EUR/USD exchange rate does not participate in the adjustment process. It is the EUR/GBP exchange rate that takes the complete burden of adjustment. After implementing the identifying restriction consistent with the no arbitrage condition that is not rejected as given by the corresponding LR-test with a p-value of 0.105, the results turn to the opposite and the EUR/GBP exchange rate is then weakly exogenous (t-statistic of 0.631 for the EUR/GBP and of -2.656 for the EUR/USD). The result is surprising because it shows that now the most traded exchange rate, namely the EUR/USD exchange rate, adjusts to a (in terms of market turnover) less important exchange rate when a disequilibrium occurs. The following economic analysis will shed light on that issue.

For the third period, with a p-value of 0.190 the identifying restriction cannot be rejected either and the less important (in terms of market turnover) exchange rate, namely the GBP/USD, takes

[45] Weak exogeneity means that a variable does not adjust to the long-run relationship. It can also be tested by applying a LR test with a restriction of the α-vector (Johansen, 1995). Since in a bivariate cointegration analysis it is expected that the results do not differ, these tests are not reported here.

the whole burden of adjustment to long-run equilibrium. In this period, the EUR/USD is weakly exgoneous. Concerning the stability tests, the results are fairly robust because no autocorrelation and autoregressive heteroskedasticity remains in the residuals as indicated by the Lagrange multiplier tests presented in tables 8 and 9. Only for the first sub-period in both the unrestricted and restricted estimation do the p-values show weak evidence of ARCH effects. This problem does not bear any weight because the Johansen test is quite robust when heteroskedasticity remains in the residuals (Rahbek, Hansen, and Dennis, 2002).

5.3 Analyzing the coincidence of economic variables

We have shown that currencies can be classified as the same asset during short periods. Referring to the above outlined economic argumentation, we shall look at the reasons why the exchange rates exhibit such a strong comovement, i.e. whether fundamental aspects are responsible for that finding. For this purpose, we need to formulate an exchange rate determination process based upon the model outlined in section 2.

Departing from the general asset approach we draw on the monetary model as outlined in section 2.1. Without losing the thought behind the model, a very strong assumption concerning the fundamentals is applied. It is assumed that all fundamental processes follow a random walk process and that market participants cannot build expectations far ahead. For this reason, the parameter b in equation (2) runs to zero, i.e. a small weight lies on the expected changes of the exchange rate and the exchange rate determination equation reduces to a similar one as given in equation (1). In a different setting this special case is treated in Engle and West (2004, 2005). Thus, regarding the fundamentals, it is assumed that the market participants focus on the currently known fundamentals (e.g. Roberts, 2005 and Poloviita, 2006 for price developments).

The monetary approach is generally applied to longer horizons usually starting with a frequency of quarterly, but sometimes also monthly, data (e.g. Goldberg and Frydman, 2001; Groen, 2002). When applying the monetary approach to higher frequencies the overshooting phenomenon due to slower adjustments of the goods markets' prices as highlighted by Dornbusch (1976) becomes relevant. Since our focus is on weekly data, i.e. choosing a frequency that can be seen as the transmission from very short-run influences on medium-term developments, we cannot smoothly apply the monetary model within an econometric analysis. In addition, the data are only available for monthly or lower frequency observations, in particular for money supply and income.

The choice of the interest rates' maturity in the monetary demand equation is not quite clear in the literature. On the one hand, it is argued that a long-term maturity should be taken because longer horizons are relevant for determining the opportunity costs of holding money. On the other hand, the same argument is applied to money market interest rates (see Nelson, 2003, for a discussion). Short-term interest rates are formed under the conditions prevailing on the credit market and also absorb temporary liquidity effects, whereas interest rates with a long-term maturity reveal anticipated returns to capital and anticipated inflation (Meltzer, 1998, p. 16).

The last point can be seen by using the term structure of interest rates; the long-term interest rate can be split up into its components.[46] It consists of the short-term interest rate and the rate of

[46]Liquidity premia are neglected because it assumed that the consumers' liquidity preferences across the countries are very similar. In addition, risk premia are also excluded because we treat yields of government bonds that are typically seen as risk-free.

inflation that is expected for the maturity (π_t^e). Hence,

$$i_{l,t} = i_{k,t} + \pi_{l,t}^e \tag{17}$$

with l as the maturity of the long-term interest rate and k as that of short-term interest rate (the prevailing money market rate). Mishkin (1990) argues that information about expected rates of inflation is more intensively embedded in long-term interest rates than in short-term ones. Consequently, equation (17) can be used to replace the domestic and foreign interest rates in equation (1). If it is accepted that longer-term interest rates solely reflect expected inflation rates, the result is equation (18), namely the real interest rate differential model (Frankel, 1979).

$$s_t^{1j} = (m_t^1 - m_t^j) + \alpha_1(y_t^1 - y_t^j) + \beta_0(i_{k,t}^1 - i_{k,t}^j) + \beta_1(i_{l,t}^1 - i_{l,t}^j) \tag{18}$$

Equation (18) is able to catch the expectations concerning the rate of inflation and can thus reveal the dynamics of the overshooting phenomenon. In the overshooting phenomenon the domestic currency appreciates if, e.g., the domestic money supply shrinks, which causes the money market rates to rise. The reason is that higher interest rates compared to the foreign ones signal the market higher profit opportunities. In contrast, the long-term interest rates decrease due to lower expected inflation rates. If the change in the long-term interest rate is fully caused by changes in expected rates of inflation the exchange rate reacts inversely to long-term interest rates. Hence, β_0 is negative and β_1 positive respectively.

A general problem with the variables in the monetary approach is that all three variables in equation (1) are interlinked and only the money supply is exogenously given. The interest rates react to changes in money supply and income, whereas income reacts to changes in interest rates. As a way out of the endogeneity in the short-run we introduce feasible assumptions.

Assuming that consumers smooth their consumption (Fisher, 1930) the demand for liquidity directed to transactions will be smoothed as well. Changes in the demand for money will be solely reflected in variations in the interest rates. Regarding changes in money supply, the consequences in the short-run will be reflected in changes in the interest rates due to liquidity effects, as well as in the expectations concerning the price level in the future. This is particularly true when regarding the overshooting of exchange rates. For this reason, the short-term interest rates and the expected inflation rates are seen as the key variables in the determination process of the asset approach.

Bearing the monetary approach in mind, we start with an inspection of the general assumption of the monetary approach before focusing on the income developments. The flex price monetary approach assumes that the PPP holds continuously, whereas the sticky prices version works with inflation rate expectations based upon the money stock. In reality, expectations concerning the inflation rates are predominantly directed to the current value of the rate of inflation and are extrapolated into the future (c.f. Roberts, 2005 for the USA, Paloviita, 2006 for the Euro area). We depart from the strict form of PPP and focus on the relative form which states that the change of exchange rates is equal to the difference in rates of inflation. The currency with a greater rate of inflation is expected to depreciate. In figure 5 the inflation rate differentials between the EMU and the USA as well as between the EMU and the UK are depicted in panel (a), whereas panel (b) shows the differentials between the USA and both the EMU and the UK. During the first period the inflation rate differential between the EMU and the UK seems to comove with the differential between the

EMU and the US, whereas the differential between the USA and the UK is close to zero at the end of this sub-period supporting the stability of the GBP/USD exchange rate. Not covered by the inflation rate differentials, the Euro has a tendency to depreciate. A similar pattern with respect to the inflation rate differentials arises for the third period. Here, the differentials between the USA and both the EMU and the UK move parallel. Again, the inflation rate differential between the EMU and the UK is close to zero at the end of this sub-period. But without an empirical investigation, which is difficult to apply because rates of inflation are only available at a monthly frequency, the linkage between differentials and exchange rate cannot be discovered properly.

Turning to the income component, a visual inspection helps discover potential comovements. For this reason, the business cycle correlations calculated on a three-year moving window are depicted in figure 6. In the first period, the economic variables in the USA and the UK should be very similar. As can be seen from the figure, the correlations between the business cycles are negative for the USA and the UK when the period of stationarity regarding the GBP/USD begins. In the following month, the three-year moving correlation increases remarkably. This is in line with the findings of the literature that the UK is relatively synchronized with the USA at the end of the 1990s (e.g. Artis, 2006). With respect to the second period, all three correlations are positive and approximately at the same level, supporting the stationarity of exchange rates between each country. For the third period, it can be considered that the business cycles' correlation between the UK and the EMU is positive. It can be seen that the synchronization of the economy is much more pronounced in the second and third period, as already highlighted by e.g. Perez, Osborn and Artis (2006).

In the following, the demand for liquidity and the money stock are summarized within the vector Σ_t. Hence, it results

$$s_t^{1j} = \gamma \Sigma_t + \beta_0 (i_{k,t}^1 - i_{k,t}^j) + \beta_1 (i_{l,t}^1 - i_{l,t}^j) \tag{19}$$

whereas the vector γ contains the elasticites with respect to money supply and income.

As discussed in section 2, two currencies can be classified as being the same asset when their fundamentals coincide. To be cointegrated, the matrix Σ_t needs to be stationary. With our reasonable assumption we see the condition as fulfilled. In addition, the short-term interest rates must be either stationary or cointegrated. The same can be said regarding the expected inflation rates. Since the expectations of inflation rates cannot be observed properly at a weekly frequency we cannot investigate this issue from a time series perspective. But we know equation (17). For this reason, we can extract the expected inflation rates from the long-term interest rates and make use of the RID model. Neglecting impact factors other than short-term interest rates and expected inflation rates, the long-term interest rates' degree of integration depends on the parameter of integration of the short-term rates and the expected inflation rates. Assuming non-stationarity of each component, cointegration of long-term interest rates can only arise if the short-term interest rates as well as the expected rates of inflation are cointegrated.

Therefore, we start with the investigation of interest rates and employ weekly short-term interest rates with a maturity of one week (also to embed the uncovered interest rate parity) and of one month, as well as long-term interest rates with a maturity of 10 years. In table 7, the unit root tests on the interest rates are given for the two periods of interest, i.e. the first and the third period. In general, the results of the three unit root tests are mixed. Hence, we see evidence in favour of non-stationarity if two of the three tests support this result. In panel (a) of table 7 all interest rates of the UK, the USA and the EMU seem to be non-stationary. The pre-condition for the cointegration analysis is fulfilled with respect to the first period. The same cannot be said for the third period. Since the US monetary policy tightens and interest rates continuously rise during the third period, a linear trend is introduced into the null hypothesis of the unit root tests for the USA. In this period, the EMU and the UK interest rates are non-stationary while the US short-term interest rate seems to be stationary around a linear trend.

Regarding the results of the unit root tests, a cointegration analysis of the US and UK interest rates in the first period and the EMU and the UK in the second can be applied by using the Johansen approach with small sample corrections. We decide to introduce a linear trend in the cointegration relationship to account for trending behaviour of the cointegration relationship and to test for its presence. The results are presented in figure 10. Indeed, the short-term interest rates of the US and the UK are cointegrated in the first period because the null hypothesis of no cointegration can be rejected with a p-value of 0.000 and 0.003 respectively. With the help of a LR-test, we test on the absence of the linear trend in the cointegration relationship (Johansen, 1995, p. 162). The LR test can reject the null hypothesis of the absence of a linear trend restricted to the cointegration relationship with a p-value of 0.046 for the maturity of one week and of 0.001 for one month. By contrast, the long-term interest rates are not cointegrated. The inspection of the estimated VECM in table 11 yields that the US interest rate with a maturity of one week is weakly exogenous (t-statistic of the α_{US} coefficient -1.12) and the UK interest rate with a maturity of one week adjusts to the US rate (t-statstic of the α_{UK} coefficient is -3.371). These results do not hold for the interest rates with a maturity of one month. Here, both interest rates are completely endogenous. In Wang, Yang and Li (2007), who investigate Eurocurrency interest rate linkages within a multivariate framework, the UK interest rate is not caused by the US interest rate during the period 1994-1998. However, for our sub-sample it can be concluded that the US interest rate does cause the UK interest rate. Returning to the results of table 8, the linkages in the interest rates' long-run relationship does not support the finding that the EUR/GBP exchange rate causes the EUR/USD exchange rate after implementing the no arbitrage condition.

The cointegration vectors for both short-term interest rates also show that there is no exact comovement of the short-term interest rates, first of all because of the cointegration parameter for the UK interest rate (0.264) and, secondly, because a linear trend prevails in the cointegration space. A linear trend in the cointegration relationship of the interest rates should generate a linear trend in the cross rate as well. Two exchange rates are cointegrated but drift apart deterministically. Since this is not the case, the market participants must have knowledge about the steadily diverging interest rates because they are not embedded in the cross rate's development. Thus, they do not care about the linear trend because they interpret it as a required adjustment in the UK or US economy's equilibrium or a change in the risk premia. Hence, the cross rate is stable.

Contrary to the first period, a similar result cannot be confirmed by investigating the interest rates of the third period. Here, the EMU and the UK interest rates are not cointegrated (panel (b) in table 10). The null hypothesis of zero rank cannot be rejected with p-values greater than of 0.400. When looking at figure 8, a strong comovement of the interest rate differentials of the EMU and the UK to the USA seems to prevail. For this reason, we apply the cointegration analysis to the interest rate differentials. The results of the mentioned unit root tests are presented in panel (b) of table 7. Since a linear trend could be present in both interest rate differentials, the unit root tests are carried out with the assumption that the linear trend is absent or present. Both tests give evidence that the interest rate differentials are non-stationary. Only the interest rate differential from the UK to the USA appears stationary.

The results of the cointegration analysis concerning the interest rate differentials are given in panel (c) of table 10. We also start with a model in which a linear trend is restricted to lie within the cointegration space. As can be seen, the null hypothesis of no cointegration can be rejected with a p-value of 0.044 for the interest rate differentials with a maturity of one week but not for the interest rate with a maturity of one month. In the first case the corresponding LR test strongly rejects the null hypothesis of no linear trend in the cointegration space which can be seen when looking at the p-value of 0.002. However, the null hypothesis is strongly rejected with a p-value of 0.314 for the second case. Here, a common linear trend in the interest rate differentials cannot be rejected. Nevertheless, the finding of no cointegration remains but with a p-value of 0.142. As the cointegration tests show, the interest rate differentials based upon interest rates with a maturity of one month share one common linear trend in levels.

By evaluating the adjustment coefficients after estimating the VECM, as given in table 12, both one week differentials are completely endogenous and adjust to each other. Although the cointegration tests cannot reject the null hypothesis of rank zero for the interest rate differential with a maturity of one month, we estimate the VECM with the assumption of rank 1 anyway. The results indicate that the cointegration parameter and the adjustment coefficient for the UK are significantly different from zero but with the curtailing that the adjustment coefficient is significantly different from zero at the 5% level. Compared with the estimation results for one week maturity, this result is more consistent because the cointegration coefficient for the UK has the correct sign and is close to zero. Although the LR-test on the restricted linear trend indicates that the setting of the model is correct, a similar result can be obtained when the model with a linear trend in levels with induced cointegration rank 1 is estimated.[47] Thus, the specification of the model is responsible for that finding. Based upon the interest rate differentials with a maturity of one month, the UK-US interest rate differential comoves with and adjusts to the EMU-US interest rate differential.

5.4 Analyzing the impact of economic variables

In order to quantify the impact of interest rates and interest rate differentials on the exchange rates, simple regressions are applied, the results are presented in table 13. To avoid the problem of spurious regressions due to the non-stationarity of the exchange rate, its change is regressed on (changes of) explanatory variables. Furthermore, we apply different economic models that take account of the discussion in the literature. In particular, the discussion about the correct interest rate in the money demand is incorporated to the extent that both the short-term (with a one-month maturity) in model I and long-term interest rates in model II are utilized. To get a deeper insight into

[47] The cointegration parameter for the UK-US differential is -0.703 with a t-value of -11.654. The adjustment coefficients are not significantly different from zero (t-values -1.459 for the EMU-US differential and 0.441 for the UK-US differential respectively.

the importance of different exchange rates the interest rates are also used as explanatory variables (models III and IV). The validity of the uncovered interest rate parity (UIP) of one week is also tested in model V with the one-period lagged short-term interest rate differential with a maturity of one week. The RID model is estimated separately in model VIII in which the short-term interest rate with a maturity of one month and the long-term interest rate are included.

Contrary to many contributions in the literature, the uncovered interest rate parity holds for the EUR/GBP exchange rate but not for the EUR/USD rate in the first period (model V). In addition, the EUR/GBP exchange rate is driven by long-term interest rates (model I and VIII). The insignificant short-term interest rates and the negative sign of the long-term interest rate differential that is inconsistent with the RID model indicate that anticipated returns on capital (as argued by Meltzer, 1998) are responsible for a small part of the exchange rate behaviour. Model III shows that changes in the UK long-term interest rates have a significant impact on the EUR/GBP exchange rate. It should be noted that also the stationary GBP/USD exchange rate is influenced by the long-term interest rate differential. Here, both long-term interest rates have a similar effect on the changes in the exchange rates. The reason can be seen in the increased synchronization of their business cycles.

Furthermore, the estimation results from table 13 show that changes in the EUR/GBP exchange rate are driven by the UK long-term interest rate and the UIP with a horizon of one week hold, whereas the UK short-term interest rate is cointegrated with and caused by the US rate. Recapitulating the results of table 8, the EUR/USD exchange rate reacts to the EUR/GBP exchange rate. Although there is evidence in favour of comovements in exchange rates led by fundamental factors, the discovered linkages cast doubt on whether money market fundamentals are solely responsible for that finding. A factor that seems to have reinforced the development has to do with the establishment of the European Monetary Union. The beginning of the second stage coincides with our start of the period of observation. During the year 1996 it became clear that the convergence criteria, fixed in the treaty of Maastricht, that had to be fulfilled for entering the third stage, in particular the fiscal ones, were not met by most of the European countries (see Mayes and El-Agraa, 2004, p. 166). Hence, the comovements of nominal exchange rates can be a reflection of the synchronization of US and UK money markets, but non-fundamental factors coming from the denomination currency seem to be important too. The EUR/GBP rate that is driven by fundamentals seems to act as a reference point for the EUR/USD rate. Since this aspect is beyond the scope of this paper, the issue should be subject of future research. From this point of view, the US dollar and the British pound have stronger linkages in fundamentals but the reason for which they can be seen as the same asset is also due to variables from the non-fundamental domain.

From table 7 it follows that the long-term interest rate differential is non-stationary while the short-term interest rate differential is stationary. Hence, the differential in expected inflation rates cannot be stationary. Regarding the stationarity of the cross rate this would mean that inflation rate expectations are not relevant in determining the exchange rate during this period. As can be seen in table 14, the change of the EUR/GBP exchange rate indeed does not react to changes in the long-term interest rate differential but to changes in money market rates consistent with the RID model, whereas the impact stems from the change in the UK money market rate. Regarding the stationary EUR/GBP exchange rate the impact of the short-term interest rates is consistent with the liquidity effect of the overshooting model but without generating considerable expectations about future inflation.

In the following, it is tested for the EUR/USD and GBP/USD rates whether the UIP holds in period three. Bearing in mind the uncovered interest rate parity, the interest rate differential can be replaced by the expected change in the exchange rate. Hence, the cointegration of interest rate differentials can basically be a reflection of a comovement in the expected returns as long as the uncovered interest rate parity holds. While the sign of the EMU-USA interest rate differential is mostly positive, the sign of the UK-USA differential is predominantly negative in the sub-sample. This also has consequences for the expected change in the exchange rates. While the Euro is expected to appreciate vis-à-vis the US-Dollar the British pound is expected to depreciate in most of the sub-sample, except the short period during which the USA has a higher interest rate level than the UK. At this, the direction of changes in the interest rate differential could be important (Evans and Lyons, 2002, p. 175) and can be seen as evidence in favour of transmitted expectations regarding future interest rates across the market. As can be seen in table 14 the UIP does not hold in all cases (model V). However, the results of the regression of the exchange rates on the (one-period lagged) change in interest rate differentials reveal the same sign and approximately the same magnitude of the regression coefficient (model V). An increase in the change of the interest rate differential increases both exchange rates by the same magnitude. A look at figure 8 shows that the period of the exchange rates' strong comovement coincides exactly with the tightening of the US monetary policy. The results from model VII in table 14 support this supposition because only the (lagged) change in the US interest rate with a maturity of one week is significant. Thus, the US monetary policy's strength seems to be responsible for the initialisation of the exchange rates' comovement by generating expectations of higher US interest rates. The channel seems to run from the EUR/USD market to the GBP/USD market because of the weak exogeneity of both the interest rate differential (interest rates with a maturity of one month) and the EUR/USD rate. The simultaneous negative effect of changes in the US interest rates with a maturity of one month is puzzling. But if it is accepted that the change in monetary policy is a response of US monetary authorities on the slowdown of the economy, the negative sign is then a reflection of the market participants' mistrust regarding the success of the stabilization's efforts. The same effect should then be found in the very short-term interest rates and if this really were the explanation, in the EUR/USD exchange rate too. But this is not the case.

6 Conclusion

We can show that stronger comovements between exchange rates exist in reality. Our results give an explanation for Sephton and Larsen's (1991) results with which they conclude that cointegration of exchange rates depends on the period of observation. At the beginning, we theoretically derive under which circumstances exchange rates can be bivariately cointegrated. This is done by applying a rational expectation approach in the vein of Mussa (1976) and Frenkel and Mussa (1980). The results of these approaches can be used to explain why cointegration is only present in short sub-samples. It requires that the fundamentals of the two countries, whereas one of whose currencies is not taken as the denomination currency, are cointegrated (if non-stationary). Hence, both currencies denominated in the same currency are truly the same asset, as introduced by Hakkio and Rush (1989) or Baffes (1994).

The results yield that periods of stronger comovements between the USD and GBP based upon the Euro prevail during the 1990s and periods of comovements between EUR and GBP prevail since the introduction of the Euro. Furthermore, no long-run relationships can be discovered across the observed exchange rates. In the first period the short-term interest rates of the USA and the UK are cointegrated. In the second period the short-term interest rates of the UK and the EMU are not cointegrated but the interest rate differentials between the EMU and the USA, and the UK, and the USA are. Our results tend to indicate that the coincidence of fundamental variables is responsible for seeing currencies as the same asset. During the two mentioned periods money market variables seem to be responsible for the stronger comovement of exchange rates and stationarity of exchange rates. However, the investigation of the fundamentals' impact on exchange rates provides evidence that factors other than money market variables played an important role during the 1990s. The conclusion can be drawn that an application of cointegration techniques to exchange rates for market efficiency considerations in short periods has to be accompanied by an investigation of fundamentals although the implementation of the auxiliary condition, the no arbitrage condition by Dwyer and Wallace (1992), holds.

This paper gives four major innovations to the literature. It first shows under which conditions exchange rates can be bivariately cointegrated. Secondly, it uses the cross-rate identity to test for cointegration, i.e. deducing recursively. Thirdly, it applies the cointegration methodology within a triangular framework by detecting cointegration between exchange rates that are not only denominated in U.S. dollar. And lastly, it shows that comovements between two exchange rates exist in a narrower sense but only in short periods, whereas the economic variables which have caused the relationship are explored.

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Data Appendix

Exchange rates	
EURO TO US \$ (WMR&DS) - EXCHANGE RATE	EUDOLLR(ER)
GERMAN MARK TO US \$ (WMR) - EXCHANGE RATE	DMARKE\$(ER)
JAPANESE YEN TO US \$ (WMR) - EXCHANGE RATE	JAPAYE\$(ER)
UK £ TO US \$ (WMR) - EXCHANGE RATE	UKDOLLR(ER)
JAPANESE YEN TO EURO (WMR) - EXCHANGE RATE	JPEURSP(ER)
UK £ TO EURO (WMR&DS) - EXCHANGE RATE	UKEURSP(ER)
GERMAN MARK TO UK £ (WMR) - EXCHANGE RATE	DMARKER(ER)
JAPANESE YEN TO UK £ (WMR) - EXCHANGE RATE	JAPAYEN(ER)
GERMAN MARK TO EURO SYNTHETIC (DS) - EXCHANGE RATE	DMEURSN(ER)
US \$ TO EURO (DS SYNTHETIC) - EXCHANGE RATE	USEURSN(ER)
UK £ TO EURO (DS SYNTHETIC) - EXCHANGE RATE	UKEURSN(ER)
JAPANESE YEN TO EURO (DS SYNTHETIC) - EXCHANGE RATE	JPEURSN(ER)

Short-term interest rates	
MNY MKT - OVERNIGHT FRANKFURT BANKS - MIDDLE RATE	BDMNYON
GERMANY INTERBANK 1 MONTH - OFFERED RATE	FIBOR1M
GERMANY EU-MARK 1WK (FT/ICAP) - MIDDLE RATE	ECWGM1W
UK INTERBANK 7 DAY - MIDDLE RATE	LDNIB7D
UK INTERBANK 1 MONTH - MIDDLE RATE	LDNIB1M
US EURO\$ DEP. 1 MTH (BID,LDN) - MIDDLE RATE	FREDD1M
US FED FUNDS EFFECT RATE (7 DAY AVG) - MIDDLE RATE	FRFEDFW
EURIBOR 1 WEEK - OFFERED RATE	EIBOR1W
EURIBOR 1 MONTH - OFFERED RATE	EIBOR1M
GERMANY EU-MARK 1WK (FT/ICAP) - MIDDLE RATE	ECWGM1W
GERMANY INTERBANK 1 MONTH - OFFERED RATE	FIBOR1M

Long-term interest rates	
BD BENCHMARK 10 YEAR DS GOVT. INDEX - RED. YIELD	BMBD10Y(RY)
EMU BENCHMARK 10 YR. DS GOVT. INDEX - RED. YIELD	BMEM10Y(RY)
UK BENCHMARK 10 YEAR DS GOVT. INDEX - RED. YIELD	BMUK10Y(RY)
US BENCHMARK 10 YEAR DS GOVT. INDEX - RED. YIELD	BMUS10Y(RY)
BD BENCHMARK 10 YEAR DS GOVT. INDEX - RED. YIELD	BMBD10Y(RY)

Consumer price index	
BD CPI NADJ	BDOCP009F
US CPI NADJ	USOCP009F
UK CPI NADJ	UKOCP009F
EK HARMONISED CPI NADJ	EJOCP049F
BD CPI NADJ	BDOCP009F

Industrial production index	
US PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION)	USOPRI35G
BD PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION)	BDOPRI35G
UK PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION)	UKOPRI35G
EM PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION)	EMOPRI35G

Datastream mnemonics in the right column.

Table 2: Unit Root Tests for the exchange rates for the whole period of observation

January 1994 till June 2007

	Levels						**First Differences**					
	Phillips-Perron Test		KPSS Test		DF-GLS (Elliot et al. 1996)		Phillips-Perron Test		KPSS Test		DF-GLS (Elliot et al. 1996)	
Exchange rates	Z(t)	p-value	lags	Test statistic	lags	DF-GLS μ	Z(t)	p-value	lags	Test statistic	lags	DF-GLS μ
EUR/USD	-1.049	(0.735)	18	0.767 **	1	-0.918	-26.892	(0.000)	22	0.221	19	-5.130 **
GBP/USD	-0.872	(0.797)	18	1.650 **	12	0.791	-25.198	(0.000)	26	0.157	19	-1.197
JPY/USD	-2.176	(0.215)	18	0.639 *	1	-2.132 *	-25.395	(0.000)	17	0.048	15	-4.656 **
GBP/EUR	-1.319	(0.620)	18	1.680 **	1	-0.765	-28.417	(0.000)	5	0.138	19	-1.679
JPY/EUR	-1.025	(0.744)	18	0.594 *	1	-0.943	-26.701	(0.000)	14	0.181	10	-6.645 **
JPY/GBP	-0.972	(0.763)	18	1.620 **	1	-0.315	-25.803	(0.000)	16	0.103	19	-3.130 **

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The p-values of the Phillips-Perron test are calculated on the basis of MacKinnon (1996) and given in parentheses. The KPSS test bases upon Kwiatkowski et al. (1992). The autovariances of the KPSS test are weighted by Bartlett kernel. The DF-GLS test bases upon Elliot et al. (1996) and the lag selection is carried out by using the modified Akaike information criterion (MAIC) as developed by Ng and Perron (2001). For the Phillips-Perron test and the DF-GLS test the null hypothesis it is assumed that the time series are non-stationary. A rejection would be evidence in favour of stationarity. In the case of the KPSS test the null hypothesis is stationarity. A non-rejection would be evidence in favour of stationarity. Critical values for the KPSS test are 0.463 for 5% and 0.347 for 1% signifiance level, for the DF-GLS test -1.95 for 5% and -2.58 for 1% significance level.

Table 3: Unit Root Tests for the exchange rates before the introduction of the Euro

16 October 1996 till 19 May 1999

	Levels						**First Differences**					
	Phillips-Perron Test		KPSS Test		DF-GLS (Elliot et al. 1996)		Phillips-Perron Test		KPSS Test		DF-GLS (Elliot et al. 1996)	
Exchange rates	Z(t)	p-value	lags	Test statistic	lags	DF-GLS μ	Z(t)	p-value	lags	Test statistic	lags	DF-GLS μ
EUR/USD	-1.354	(0.604)	11	1.090 **	1	-0.856	-16.481	(0.000)	14	0.284	7	-4.831 **
GBP/USD	-2.230	(0.196)	11	1.560 **	1	-0.588	-16.474	(0.000)	12	0.045	14	-1.092
JPY/USD	-1.191	(0.678)	11	1.700 **	1	-1.165	-15.245	(0.000)	10	0.144	13	-3.059 **
GBP/EUR	-0.587	(0.874)	11	1.490 **	3	-0.533	-16.975	(0.000)	17	0.300	15	-1.395
JPY/EUR	-1.816	(0.372)	11	1.360 **	1	-1.442	-14.021	(0.000)	10	0.062	8	-3.800 **
JPY/GBP	-0.921	(0.781)	11	1.900 **	3	-0.910	-14.788	(0.000)	11	0.135	13	-2.347 *

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The p-values of the Phillips-Perron test are calculated on the basis of MacKinnon (1996) and given in parentheses. The KPSS test bases upon Kwiatkowski et al. (1992). The autovariances of the KPSS test are weighted by Bartlett kernel. The DF-GLS test bases upon Elliot et al. (1996) and the lag selection is carried out by using the modified Akaike information criterion (MAIC) as developed by Ng and Perron (2001). For the Phillips-Perron test and the DF-GLS test the null hypothesis it is assumed that the time series are non-stationary. A rejection would be evidence in favour of stationarity. In the case of the KPSS test the null hypothesis is stationarity. A non-rejection would be evidence in favour of stationarity. Critical values for the KPSS test are 0.463 for 5% and 0.347 for 1% signifiance level, for the DF-GLS test -1.95 for 5% and -2.58 for 1% significance level.

Table 4: Unit Root Tests for the exchange rates after the introduction of the Euro

16 June 2004 till 14 June 2006

	Levels						**First Differences**					
	Phillips-Perron Test		KPSS Test		DF-GLS (Elliot et al. 1996)		Phillips-Perron Test		KPSS Test		DF-GLS (Elliot et al. 1996)	
Exchange rates	Z(t)	p-value	lags	Test statistic	lags	DF-GLS μ	Z(t)	p-value	lags	Test statistic	lags	DF-GLS μ
EUR/USD	-0.546	(0.883)	13	2.330 **	1	-0.710	-21.229	(0.000)	24	0.407	17	-2.985 **
GBP/USD	-0.378	(0.914)	13	2.350 **	3	-0.445	-19.304	(0.000)	42	0.310	1	-14.097 **
JPY/USD	-2.130	(0.233)	13	0.251	1	-2.078 *	-20.443	(0.000)	23	0.075	15	-4.275 **
GBP/EUR	-2.042	(0.268)	13	1.680 **	1	-1.327	-22.998	(0.000)	12	0.179	17	-1.695
JPY/EUR	-0.279	(0.928)	13	2.440 **	1	-0.564	-22.895	(0.000)	11	0.473 *	17	-3.336 **
JPY/GBP	-0.227	(0.935)	13	2.420 **	1	-0.493	-21.703	(0.000)	15	0.361	15	-3.968 **

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The p-values of the Phillips-Perron test are calculated on the basis of MacKinnon (1996) and given in parentheses. The KPSS test bases upon Kwiatkowski et al. (1992). The autovariances of the KPSS test are weighted by Bartlett kernel. The DF-GLS test bases upon Elliot et al. (1996) and the lag selection is carried out by using the modified Akaike information criterion (MAIC) as developed by Ng and Perron (2001). For the Phillips-Perron test and the DF-GLS test the null hypothesis it is assumed that the time series are non-stationary. A rejection would be evidence in favour of stationarity. In the case of the KPSS test the null hypothesis is stationarity. A non-rejection would be evidence in favour of stationarity. Critical values for the KPSS test are 0.463 for 5% and 0.347 for 1% signifiance level, for the DF-GLS test -1.95 for 5% and -2.58 for 1% significance level.

Table 5: Log Periodogram Regression by Phillips (2007) on levels

Unit Root Log Periodogram Regression (Phillips, 2007)									
Power = 0.5	January 1994 till June 2007			16 October 1996 till 19 May 1999			16 June 2004 till 14 June 2006		
	$\hat{d}$	p(H0: d=0)	p(H0: d=1)	$\hat{d}$	p(H0: d=0)	p(H0: d=1)	$\hat{d}$	p(H0: d=0)	p(H0: d=1)
Exchange rate	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
EUR/USD	1.070	(0.000)	(0.578)	0.910	(0.003)	(0.574)	1.178	(0.000)	(0.204)
GBP/USD	1.208	(0.000)	(0.097)	0.454	(0.009)	(0.001)	0.976	(0.000)	(0.863)
JPY/USD	0.802	(0.000)	(0.116)	0.988	(0.001)	(0.941)	1.203	(0.000)	(0.145)
GBP/EUR	1.132	(0.000)	(0.294)	0.975	(0.000)	(0.876)	0.854	(0.000)	(0.297)
JPY/EUR	0.978	(0.000)	(0.863)	1.034	(0.000)	(0.833)	1.044	(0.000)	(0.753)
JPY/GBP	1.034	(0.000)	(0.787)	0.866	(0.000)	(0.404)	1.107	(0.000)	(0.446)

$\hat{d}$ is the integration parameter and shows the number a time series must be differenced to obtain stationarity.
The parameter $\hat{d}$ is the estimate of the least squares regression $\log(I_X(\lambda_s)) = \hat{c} - \hat{d}\log\left|1 - e^{i\lambda_s}{}^2\right| + resid$. P-values in parentheses.

Table 6: Johansen Cointegration Analysis for the validity of the no arbitrage condition

Johansen Cointegration Analysis						
lags	Cointegration rank	eigenvalue	λ_{trace}	p-value	AR(5)	p-value
			EUR/USD, GBP/USD and EUR/GBP			
1	r=0	0.446	421.037	(0.000)	5.869	(0.753)
	r≤1	0.005	6.030	(0.941)		
	r≤2	0.003	2.220	(0.734)		
			EUR/USD, JPY/USD and EUR/JPY			
1	r=0	0.468	449.289	(0.000)	8.047	(0.529)
	r≤1	0.006	5.379	(0.965)		
	r≤2	0.002	1.095	(0.924)		
			GBP/USD, JPY/USD and GBP/JPY			
1	r=0	0.392	355.750	(0.000)	10.663	(0.300)
	r≤1	0.007	6.378	(0.925)		
	r≤2	0.003	1.787	(0.813)		
			EUR/GBP, EUR/JPY and GBP/JPY			
1	r=0	0.459	434.32	(0.000)	10.481	(0.313)
	r≤1	0.004	4.945	(0.976)		
	r≤2	0.003	1.983	(0.778)		
	No. Of obs.	704				

λ_{trace} is the trace statistic (Johansen, 1991) corrected for small samples (Johansen, 2002). The p-values base upon Doornik (1998) and are approximated by a Gamma distribution. In the model a constant term is restricted and lies within the cointegration space. AR(p) describes a Lagrange multiplier test that tests for serial correlation in the residuals upto p lags. The LR-tests are distributed asχ^2 with nine degree of freedom.Here, the corresponding p-values are in brakets. Regarding the rank tests, the p-values base upon Doornik (1998) and are approximated by a Gamma distribution.

Table 7: Unit Root tests for interest rates

Panel (a)	Interest Rates - Levels													
	16 October 1996 till 19 May 1999							16 June 2004 till 14 June 2006						
	Phillips-Perron Test		KPSS Test			DF-GLS (Elliot et al. 1996)		Phillips-Perron Test		KPSS Test			DF-GLS (Elliot et al. 1996)	
	Z(t)	p-value	lags	Test statistic		lags	DF-GLS μ	Z(t)	p-value	lags	Test statistic		lags	DF-GLS μ
							EMU							
1w	-3.573	(0.006)	7	0.348		9	-0.804	1.531	(0.998)	7	0.994	**	3	1.340
1m	-0.033	(0.921)	7	0.327		1	-0.529	2.124	(0.999)	7	1.030	**	1	1.495
10y	-0.763	(0.830)	7	1.710	**	3	0.291	-1.999	(0.287)	7	0.551	*	1	-0.564
							UK							
1w	-0.256	(0.932)	7	0.407		5	-0.706	-3.316	(0.014)	7	0.615	*	5	-1.207
1m	-0.185	(0.940)	7	0.398		12	-1.265	-1.851	(0.356)	7	0.845	**	1	-1.494
10y	-0.639	(0.862)	7	1.740	**	6	1.163	-2.399	(0.142)	7	0.895	**	1	-0.475
							USA							
1w	-2.196	(0.208)	7	0.806	**	12	-0.811	-6.194	(0.000)	3	0.131		10	-1.224
1m	-1.346	(0.608)	7	0.807	**	4	-0.789	-3.372	(0.055)	6	0.075		4	-2.179
10y	-1.431	(0.568)	7	1.540	**	1	-0.559	-1.346	(0.608)	7	0.701	*	1	1.198

Panel (b)	Interest Rates differentials to the USA - 16 June 2004 till 14 June 2006													
	no linear trend							linear trend						
	Phillips-Perron Test		KPSS Test			DF-GLS (Elliot et al. 1996)		Phillips-Perron Test		KPSS Test			DF-GLS (Elliot et al. 1996)	
	Z(t)	p-value	lags	Test statistic		lags	DF-GLS μ	Z(t)	p-value	lags	Test statistic		lags	DF-GLS μ
							EMU							
1w	-1.933	(0.317)	7	1.410	**	9	0.62	-1.764	(0.722)	6	0.352	**	5	-0.361
1m	-2.634	(0.086)	7	1.400	**	12	0.003	-0.310	(0.989)	6	0.348	**	4	-0.227
10y	-0.758	(0.831)	7	1.240	**	1	-0.383	-2.535	(0.031)	7	0.239	**	1	-2.600
							UK							
1w	0.252	(0.975)	7	1.410	**	11	-0.209	-4.699	(0.001)	6	0.185	*	5	-1.578
1m	0.282	(0.977)	7	1.410	**	10	-0.036	-0.274	(0.222)	7	0.166	*	1	-1.724
10y	-0.099	(0.950)	7	1.340	**	1	0.249	-4.011	(0.009)	6	0.123		10	-1.647

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The p-values of the Phillips-Perron test are calculated on the basis of MacKinnon (1996) and given in parentheses. The KPSS test bases upon Kwiatkowski et al. (1992). The autovariances of the KPSS test are weighted by Bartlett kernel. The DF-GLS test bases upon Elliot et al. (1996) and the lag selection is carried out by using the modified Akaike information criterion (MAIC) as developed by Ng and Perron (2002). For the Phillips-Perron test and the DF-GLS test the null hypothesis it is assumed that the time series are non-stationary for panel (a) and the left side of panel (b). In panel (a), only in the case of the USA for the second period the null hypothesis in all cases is adjusted by taking a linear trend term into account. A rejection would be evidence in favour of stationarity around a linear trend. In the case of the KPSS test the null hypothesis is stationarity. A non-rejection would be evidence in favour of stationarity in panel (a) and in favour of stationarity around a linear trend in panel (b). Critical values (level stationarity) for the KPSS test are 0.463 for 5% and 0.347 for 1% signifiance level, for the DF-GLS test -1.95 for 5% and -2.58 for 1% significance level; for trend stationarity 0.146 for 5% and 0.216 for 1% concerning the KPSS test and -3.025 for 5% and -3.574 for 1% concerning the DF-GLS test.

Table 8: Cointegration Analysis and Estimation of the VECM for the sub-period 16 October 1996 till 19 May 1999 on the exchange rates EUR/USD and EUR/GBP

EUR/USD and EUR/GBP: 16 October 1996 till 19 May 1999				
	test statistic	p- value	test statistic	p-value
$\lambda_{trace} : r = 0$	27.588	(0.003)		
$\lambda_{trace} : r \leqslant 1$	9.378	(0.045)		
$LR(\beta = (1, -1)) : \chi^2(1)$			2.633	(0.105)
Estimation of VECM				
	unrestricted VECM		*restricted VECM*	
	coefficients	t-statistic	coefficients	t-statistic
$\beta_{EUR/USD}$	1		1	
$\beta_{EUR/GBP}$	-1.205**	[-13.004]**	-1	
μ	-0.576**	[-16.451]**	-0.496**	[-117.694]**
$\alpha_{EUR/USD}$	-0.068	[-1.368]	-0.154**	[-2.656]**
$\alpha_{EUR/GBP}$	0.109*	[2.268]*	0.037	[0.631]
Stability analysis				
	test statistic	p-value	test statistic	p-value
$AR(1) : \chi^2(4)$	3.312	(0.507)	3.374	(0.497)
$AR(5) : \chi^2(4)$	3.414	(0.491)	2.775	(0.596)
$ARCH(1) : \chi^2(9)$	14.879	(0.094)	13.43	(0.144)
$ARCH(5) : \chi^2(45)$	61.563	(0.051)	59.542	(0.072)

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The number of lags included within the VAR is 1 and the number of observations 136. λ_{trace} is the trace statistic (Johansen, 1991) corrected for small samples (Johansen, 2002). The p-values base upon Doornik (1998) and are approximated by a Gamma distribution. β refers to the coefficient in the cointegration vector and μ to the non-zero mean included in the cointegration space. α denotes the adjustment coefficients in the VECM. $LR(\beta = (1, -1))$ is a LR test statistic on a restriction of the VECM with respect to the cointegration vector. AR(p) and ARCH(p) describe Lagrange multiplier tests that test for serial correlation and ARCH effects in the residuals up to p lags respectively. t-statistics are surrounded by squared brackets and and p-values by parentheses.

Table 9: Cointegration Analysis and Estimation of the VECM for the sub-period 16 June 2004 till 14 June 2006 on the exchange rates EUR/USD and GBP/USD

EUR/USD and GBP/USD: 16 June 2004 till 14 June 2006				
	test statistic	p- value	test statistic	p- value
$\lambda_{trace} : r = 0$	20.92	(0.039)		
$\lambda_{trace} : r \leqslant 1$	4.375	(0.371)		
$LR(\beta = (1, -1)) : \chi^2(1)$			1.718	(0.190)
		Estimation of VECM		
	unrestricted VECM		*restricted VECM*	
	coefficients	t-statistic	coefficients	t-statistic
$\beta_{EUR/USD}$	1		1	
$\beta_{GBP/USD}$	-1.166**	[10.982]	-1	
μ	-0.478**	[-7.529]	-0.378**	[102.681]
$\alpha_{EUR/USD}$	0.087	[1.12]	0.025	[-0.326]
$\alpha_{GBP/USD}$	0.245**	[3.371]	0.195**	[2.637]
		Stability analysis		
	test statistic	p-value	test statistic	p-value
$AR(1) : \chi^2(4)$	3.663	(0.454)	2.423	(0.658)
$AR(5) : \chi^2(4)$	9.739	(0.418)	3.764	(0.439)
$ARCH(1) : \chi^2(9)$	4.972	(0.837)	5.202	(0.816)
$ARCH(5) : \chi^2(45)$	45.652	(0.445)	46.801	(0.398)

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The number of lags included within the VAR is 1 and the number of observations 136. λ_{trace} is the trace statistic (Johansen, 1991) corrected for small samples (Johansen, 2002). The p-values base upon Doornik (1998) and are approximated by a Gamma distribution. β refers to the coefficient in the cointegration vector and μ to the non-zero mean included in the cointegration space. α denotes the adjustment coefficients in the VECM. $LR(\beta = (1, -1))$ is a LR test statistic on a restriction of the VECM with respect to the cointegration vector. AR(p) and ARCH(p) describe Lagrange multiplier tests that test for serial correlation and ARCH effects in the residuals up to p lags respectively. t-statistics are surrounded by squared brackets and and p-values by parentheses.

Table 10: Cointegration Analysis of interest rates and interest rates differentials for subperiods

Panel (a)	USA and UK Interest Rates: 16 October 1996 till 19 May 1999							
	lags	Cointegration rank	eigenvalue	λ_{trace}	p-value	LR on restricted trend in CIV	AR(1)	ARCH(1)
restricted linear trend								
1 week	4	**r=0**	0.265	43.528	(0.000)	3.976	4.093	12.062
		r≤1	0.034	3.508	(0.805)	(0.046)	(0.325)	(0.210)
1 month	5	**r=0**	0.187	33.964	(0.003)	11.314	7.474	39.92
		r≤1	0.051	6.853	(0.370)	(0.001)	(0.113)	(0.000)
10 years	1	r=0	0.107	18.156	(0.341)	10.464	5.16	2.574
		r≤1	0.021	2.916	(0.875)	(0.001)	(0.271)	(0.979)
No. Of obs.	136							
Panel (b)	EMU and UK Interest Rates: 16 June 2004 till 14 June 2006							
	lags	Cointegration rank	eigenvalue	λ_{trace}	p-value	LR on restricted trend in CIV	AR(1)	ARCH(1)
restricted linear trend								
1 week	2	r=0	0.104	16.684	(0.447)	4.128	1.207	9.97
		r≤1	0.055	5.194	(0.577)	(0.042)	(0.877)	(0.353)
1 month	2	r=0	0.125	16.919	(0.429)	8.544	5.186	43.271
		r≤1	0.039	3.701	(0.781)	(0.003)	(0.269)	(0.000)
10 years	1	r=0	0.089	13.511	(0.701)	0.277	1.189	5.316
		r≤1	0.037	3.929	(0.752)	(0.598)	(0.880)	(0.806)
unrestricted constant						LR on linear trend in levels		
10 years	1	r=0	0.087	10.688	(0.235)	0.320		
		r≤1	0.013	1.357	(0.244)	(0.572)		
restricted constant						LR on constant in CIV		
10 years	1	r=0	0.091	11.485	(0.504)	7.935		
		r≤1	0.016	1.657	(0.836)	(0.005)		
No. Of obs.	105							
Panel (c)	EMU and UK Interest Rates Differentials to the USA: 16 June 2004 till 14 June 2006							
	lags	Cointegration rank	eigenvalue	λ_{trace}	p-value	LR on restricted trend in CIV	AR(1)	ARCH(1)
restricted linear trend								
1 week	1	**r=0**	0.160	26.170	(0.044)	9.731	5.509	-0.722
		r≤1	0.076	8.234	(0.240)	(0.002)	(0.239)	(1.000)
1 month	2	r≤0	0.109	17.021	(0.421)	1.013	4.818	15.577
		r≤1	0.055	5.172	(0.580)	(0.314)	(0.307)	(0.076)
10 years	1	r≤0	0.154	20.668	(0.197)	13.900	7.712	7.124
		r=1	0.033	3.454	(0.812)	(0.000)	(0.103)	(0.624)
unrestricted constant						LR on linear trend in levels		
1 month	2	r≤0	0.100	12.351	(0.142)	6.969		
		r=1	0.019	1.65	(0.199)	(0.008)		
No. Of obs.	105							

λ_{trace} is the trace statistic (Johansen, 1991) corrected for small samples (Johansen, 2002). The p-values base upon Doornik (1998) and are approximated by a Gamma distribution. AR(p) and ARCH(p) describe Lagrange multiplier tests that test for serial correlation and ARCH effects in the residuals up to p lags. LR on restricted trend in CIV tests on the absence of a linear trend within the cointegration space. The null hypothesis is the absence of the linear trend. LR on linear trend in levels tests on the absence of a common linear trend in levels. The null hypothesis is the absence of the common linear trend. LR on constant in CIV tests on the absence of a constant term restricted to the cointegration space. The null hypothesis is the absence of the constant term. All LR-tests are distributed asχ^2 with one degree of freedom.

Table 11: Estimation of the VECM for the sub-period 16 October 1996 till 19 May 1999 on interest rates

Interest Rates from the USA and the UK: 16 October 1996 till 19 May 1999				
	Short term interest rates with a maturity of 1 week		Short term interest rates with a maturity of 1 month	
	coefficients	t-statistics	coefficients	t-statistics
β_{USA}	1		1	
β_{UK}	-0.264**	[-7.768]	-0.254**	[11.095]
trend	0.002**	[3.045]	0.003**	[6.143]
α_{USA}	0.087	[1.12]	-0.343**	[-4.253]
α_{UK}	0.245**	[3.371]	0.214**	[2.509]
	test statistic	p-value	test statistic	p-value
$AR(1) : \chi^2(4)$	3.642	(0.462)	1.207	(0.877)
$AR(5) : \chi^2(4)$	5.854	(0.210)	8.049	(0.090)
$ARCH(1) : \chi^2(9)$	14.453	(0.107)	9.97	(0.353)
$ARCH(5) : \chi^2(45)$	57.47	(0.101)	42.717	(0.569)

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The number of lags included within the VAR and the number of observations are given in the cointegration analysis. β refers to the coefficient in the cointegration vector and *trend* to the non-zero mean included in the cointegration space. α denotes the adjustment coefficients in the VECM. AR(p) and ARCH(p) describe Lagrange multiplier tests that test for serial correlation and ARCH effects in the residuals up to p lags respectively. t-statistics are surrounded by squared brackets and and p-values by parentheses.

Table 12: Estimation of the VECM for the sub-period 16 June 2004 till 14 June 2006 on interest rate differentials

Interest Rate Differentials from EMU and the UK to the USA: 16 June 2004 till 14 June 2006				
	Short term interest rates with a maturity of 1 week		Short term interest rates with a maturity of 1 month	
	coefficients	t-statistics	coefficients	t-statistics
β_{EMU}	1		1	
β_{UK}	5.114**	[4.153]	-0.705**	[15.793]
trend	0.243**	[4.815]		
α_{EMU}	-0.043**	[-3.242]	-0.020	[-1.125]
α_{UK}	-0.056**	[4.448]	0.058*	[2.219]
	test statistic	p-value	test statistic	p-value
$AR(1) : \chi^2(4)$	4.783	(0.310)	4.758	(0.313)
$AR(5) : \chi^2(4)$	1.356	(0.852)	0.875	(0.928)
$ARCH(1) : \chi^2(9)$	-0.105	(1.000)	21.104	(0.000)
$ARCH(5) : \chi^2(45)$	78.83	(0.001)	71.118	(0.008)

* (**) denotes rejection of the null hypothesis at the 5% (1%) significance level. The number of lags included within the VAR and the number of observations are given in the cointegration analysis. β refers to the coefficient in the cointegration vector and *trend* to the non-zero mean included in the cointegration space. α denotes the adjustment coefficients in the VECM. AR(p) and ARCH(p) describe Lagrange multiplier tests that test for serial correlation and ARCH effects in the residuals up to p lags respectively. t-statistics are surrounded by squared brackets and and p-values by parentheses.

Table 13: Impact of fundamental variables on exchange rates from 16 October 1996 till 19 May 1999 - OLS regressions

	Δs_t^{12}	constant	$\Delta(i^2_{l,t} - i^1_{l,t})$	$\Delta(i^2_{1m,t} - i^1_{1ml,t})$	$\Delta i^2_{l,t}$	$\Delta i^1_{l,t}$	Δi^2_{1mt}	$\Delta i^1_{1m,t}$	$(i^2_{1wt-1} - i^1_{1wt-1})$	$\Delta(i^2_{1w,t-1} - i^1_{1w,t-1})$	$\Delta i^2_{1w,t-1}$	$\Delta i^1_{1w,t-1}$	R^2 (centered)	DW
	$\Delta(EUR/USD)_t$	0.001	-0.016										0.014	2.119
		(0.221)	(0.217)											
I	$\Delta(EUR/GBP)_t$	**0.002**	**-0.023**										0.028	2.156
		(0.082)	(0.036)											
	$\Delta(GBP/USD)_t$	0.000	**-0.021**										0.052	2.030
		(0.569)	(0.013)											
	$\Delta(EUR/USD)_t$	0.001		0.016									0.013	2.022
		(0.179)		(0.164)										
II	$\Delta(EUR/GBP)_t$	0.002		-0.013									0.013	2.147
		(0.106)		(0.128)										
	$\Delta(GBP/USD)_t$	0.000		-0.003									0.002	1.956
		(0.774)		(0.619)										
	$\Delta(EUR/USD)_t$	0.001			-0.008	0.017							0.019	2.120
		(0.186)			(0.640)	(0.178)								
III	$\Delta(EUR/GBP)_t$	**0.002**			-0.017	**0.023**							0.030	2.142
		(0.070)			(0.276)	(0.038)								
	$\Delta(GBP/USD)_t$	0.000			**-0.195**	**0.226**							0.053	2.045
		(0.603)			(0.028)	(0.020)								
	$\Delta(EUR/USD)_t$	0.001					0.001	-0.016					0.013	2.022
		(0.185)					(0.271)	(0.1814)						
IV	$\Delta(EUR/GBP)_t$	0.002					-0.016	0.012					0.013	2.146
		(0.115)					(0.264)	(0.182)						
	$\Delta(GBP/USD)_t$	0.000					-0.011	-0.004					0.012	1.970
		(0.730)					(0.249)	(0.625)						
	$\Delta(EUR/USD)_t$	0.005							0.002				0.002	2.043
		(0.472)							(0.617)					
V	$\Delta(EUR/GBP)_t$	**0.018**							**0.005**				0.052	2.238
		(0.001)							(0.004)					
	$\Delta(GBP/USD)_t$	-0.002							0.001				0.003	1.959
		(0.493)							(0.499)					
	$\Delta(EUR/USD)_t$	0.001								0.000			0.000	2.039
		(0.185)								(0.944)				
VI	$\Delta(EUR/GBP)_t$	0.002								0.003			0.003	2.126
		(0.108)								(0.502)				
	$\Delta(GBP/USD)_t$	0.000								0.002			0.001	1.954
		(0.774)								(0.623)				
	$\Delta(EUR/USD)_t$	0.001									0.000	-0.001	0.000	2.041
		(0.187)									(0.921)	(0.852)		
VII	$\Delta(EUR/GBP)_t$	**0.002**									0.006	0.004	0.013	2.080
		(0.099)									(0.140)	(0.562)		
	$\Delta(GBP/USD)_t$	0.000									0.004	0.000	0.003	1.954
		(0.787)									(0.432)	(0.979)		
	$\Delta(EUR/USD)_t$	0.001	-0.017	0.018									0.029	2.104
		(0.216)	(0.179)	(0.090)										
VIII	$\Delta(EUR/GBP)_t$	**0.002**	**-0.022**	-0.011									0.037	2.173
		(0.082)	(0.042)	(0.174)										
	$\Delta(GBP/USD)_t$	0.000	**-0.021**	0.000									0.052	2.030
		(0.568)	(0.014)	(0.979)										

The exchange rate s_t^{12} is currency 1 expressed in currency 2. With $i_{m,t}$ as interest rates at time t and maturity m, whereas l for long-term and k for short-term interest rates, the models are: I: $\Delta s_t = \alpha + \beta_0 \Delta(i_{l,t} - i^a_{l,t})$ monetary approach: long-term interest rates, II: $\Delta s_t = \alpha + \beta_0 \Delta(i_{1m,t} - i^a_{1m,t})$ monetary approach: short-term interest rates, III: $\Delta s_t = \alpha + \beta_0 \Delta i_{l,t} - \beta_1 \Delta i^a_{l,t}$, IV: $\Delta s_t = \alpha + \beta_0 \Delta i_{1m,t} - \beta_1 \Delta i^a_{1m,t}$, V: $\Delta s_t = \alpha + \beta_0 (i_{1w,t-1} - i^a_{1w,t-1})$ uncovered interest rate parity condition, VI: $\Delta s_t = \alpha + \beta_0 \Delta(i_{1w,t-1} - i^a_{1w,t-1})$ microstructure approach (Evans/ Lyons, 2002), VII: $\Delta s_t = \alpha + \beta_0 \Delta i_{1w,t-1} - \beta_1 \Delta i^a_{1w,t-1}$, VIII: $\Delta s_t = \alpha + \beta_0 \Delta(i_{1m,t} - i^a_{1m,t}) + \beta \Delta(i_{l,t} - i^a_{l,t})$ real interest rate differential (Frankel, 1979), DW is the Durbin-Watson statistic for autocorrelation. P-values are in brackets.

Table 14: Impact of fundamental variables on exchange rates from 16 June 2004 till 14 June 2006 - OLS regressions

	Δs_t^{12}	constant	$\Delta(i_{l,t}^2 - i_{l,t}^1)$	$\Delta(i_{1m,t}^2 - i_{1ml,t}^1)$	$\Delta i_{l,t}^2$	$\Delta i_{l,t}^1$	Δi_{1mt}^2	$\Delta i_{1m,t}^1$	$(i_{1wt-1}^2 - i_{1wt-1}^1)$	$\Delta(i_{1w,t-1}^2 - i_{1w,t-1}^1)$	$\Delta i_{1w,t-1}^2$	$\Delta i_{1w,t-1}^1$	R^2 (centered)	DW
	$\Delta(EUR/USD)_t$	0.000	0.007										0.002	2.070
		(0.757)	(0.682)											
I	$\Delta(EUR/GBP)_t$	0.000	-0.146										0.007	2.082
		(0.707)	(0.429)											
	$\Delta(GBP/USD)_t$	0.000	-0.001										0.000	1.827
		(0.924)	(0.942)											
	$\Delta(EUR/USD)_t$	-0.001		0.009									0.000	2.082
		(0.931)		(0.770)										
II	$\Delta(EUR/GBP)_t$	0.000		**-0.029**									0.032	1.924
		(0.867)		(0.038)										
	$\Delta(GBP/USD)_t$	0.000		-0.002									0.000	1.818
		(0.925)		(0.945)										
	$\Delta(EUR/USD)_t$	0.000			0.002	-0.010							0.003	2.082
		(0.747)			(0.937)	(0.659)								
III	$\Delta(EUR/GBP)_t$	0.000			-0.016	0.013							0.007	2.083
		(0.694)			(0.381)	(0.510)								
	$\Delta(GBP/USD)_t$	0.000			-0.007	0.001							0.002	1.848
		(0.897)			(0.726)	(0.964)								
	$\Delta(EUR/USD)_t$	0.000					-0.022	-0.025					0.007	2.085
		(0.728)					(0.570)	(0.516)						
IV	$\Delta(EUR/GBP)_t$	0.000					-0.026	**0.030**					0.032	1.924
		(0.847)					(0.339)	(0.043)						
	$\Delta(GBP/USD)_t$	0.003					-0.039	**-0.085**					0.081	1.865
		(0.128)					(0.177)	(0.028)						
	$\Delta(EUR/USD)_t$	-0.001							-0.001				0.003	2.072
		(0.515)							(0.549)					
V	$\Delta(EUR/GBP)_t$	0.003							0.001				0.002	2.022
		(0.689)							(0.654)					
	$\Delta(GBP/USD)_t$	0.000							0.000				0.001	1.827
		(0.794)							(0.769)					
	$\Delta(EUR/USD)_t$	-0.002								**-0.037**			0.074	2.094
		(0.176)								(0.009)				
VI	$\Delta(EUR/GBP)_t$	0.000								0.018			0.019	2.025
		(0.634)								(0.147)				
	$\Delta(GBP/USD)_t$	-0.001								**-0.035**			0.078	1.798
		(0.236)								(0.007)				
	$\Delta(EUR/USD)_t$	-0.002									-0.044	**0.035**	0.075	2.102
		(0.207)									(0.1298)	(0.018)		
VII	$\Delta(EUR/GBP)_t$	0.000									0.003	-0.258	0.030	2.110
		(0.720)									(0.828)	(0.082)		
	$\Delta(GBP/USD)_t$	-0.002									-0.006	**0.047**	0.103	1.829
		(0.101)									(0.761)	(0.001)		
	$\Delta(EUR/USD)_t$	0.000	0.006	0.008									0.002	2.082
		(0.942)	(0.691)	(0.789)										
VIII	$\Delta(EUR/GBP)_t$	0.000	-0.167	**-0.030**									0.041	1.981
		(0.911)	(0.340)	(0.036)										
	$\Delta(GBP/USD)_t$	0.000	-0.001	-0.002									0.000	1.820
		(0.920)	(0.939)	(0.943)										

The exchange rate s_t^{12} is currency 1 expressed in currency 2. With $i_{m,t}$ as interest rates at time t and maturity m, whereas l for long-term and k for short-term interest rates, the models are: I: $\Delta s_t = \alpha + \beta_0 \Delta(i_{l,t} - i_{l,t}^a)$ monetary approach: long-term interest rates, II: $\Delta s_t = \alpha + \beta_0 \Delta(i_{1m,t} - i_{1m,t}^a)$ monetary approach: short-term interest rates, III: $\Delta s_t = \alpha + \beta_0 \Delta i_{l,t} - \beta_1 \Delta i_{l,t}^a$, IV: $\Delta s_t = \alpha + \beta_0 \Delta i_{1m,t} - \beta_1 \Delta i_{1m,t}^a$, V: $\Delta s_t = \alpha + \beta_0 (i_{1w,t-1} - i_{1w,t-1}^a)$ uncovered interest rate parity condition, VI: $\Delta s_t = \alpha + \beta_0 \Delta(i_{1w,t-1} - i_{1w,t-1}^a)$ microstructure approach (Evans/ Lyons, 2002), VII: $\Delta s_t = \alpha + \beta_0 \Delta i_{1w,t-1} - \beta_1 \Delta i_{1w,t-1}^a$, VIII: $\Delta s_t = \alpha + \beta_0 \Delta(i_{1m,t} - i_{1m,t}^a) + \beta \Delta(i_{l,t} - i_{l,t}^a)$ real interest rate differential (Frankel, 1979), DW is the Durbin-Watson statistic for autocorrelation. P-values are in brackets.

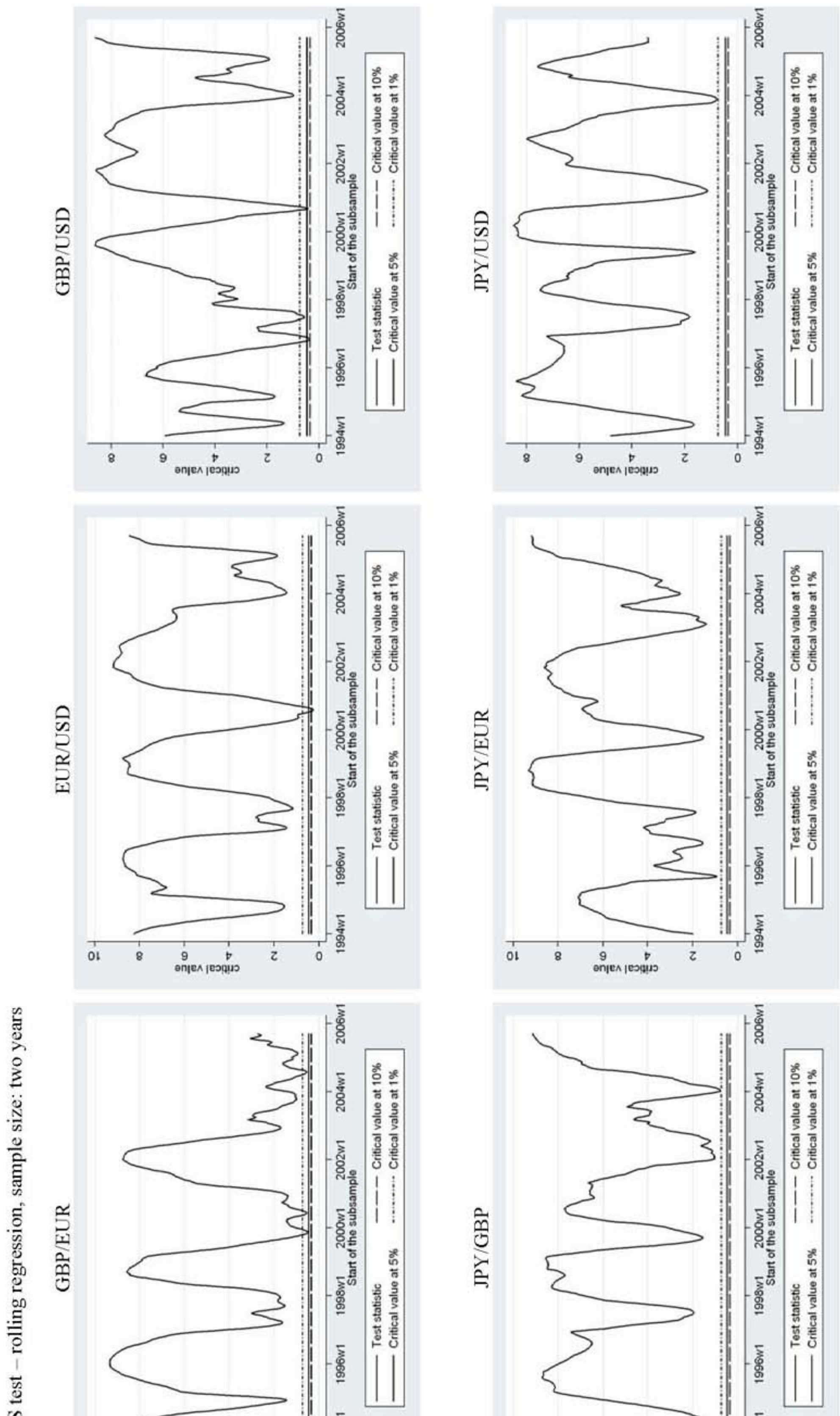

Figure 1: KPSS test - rolling regression with a fixed sample size of two years

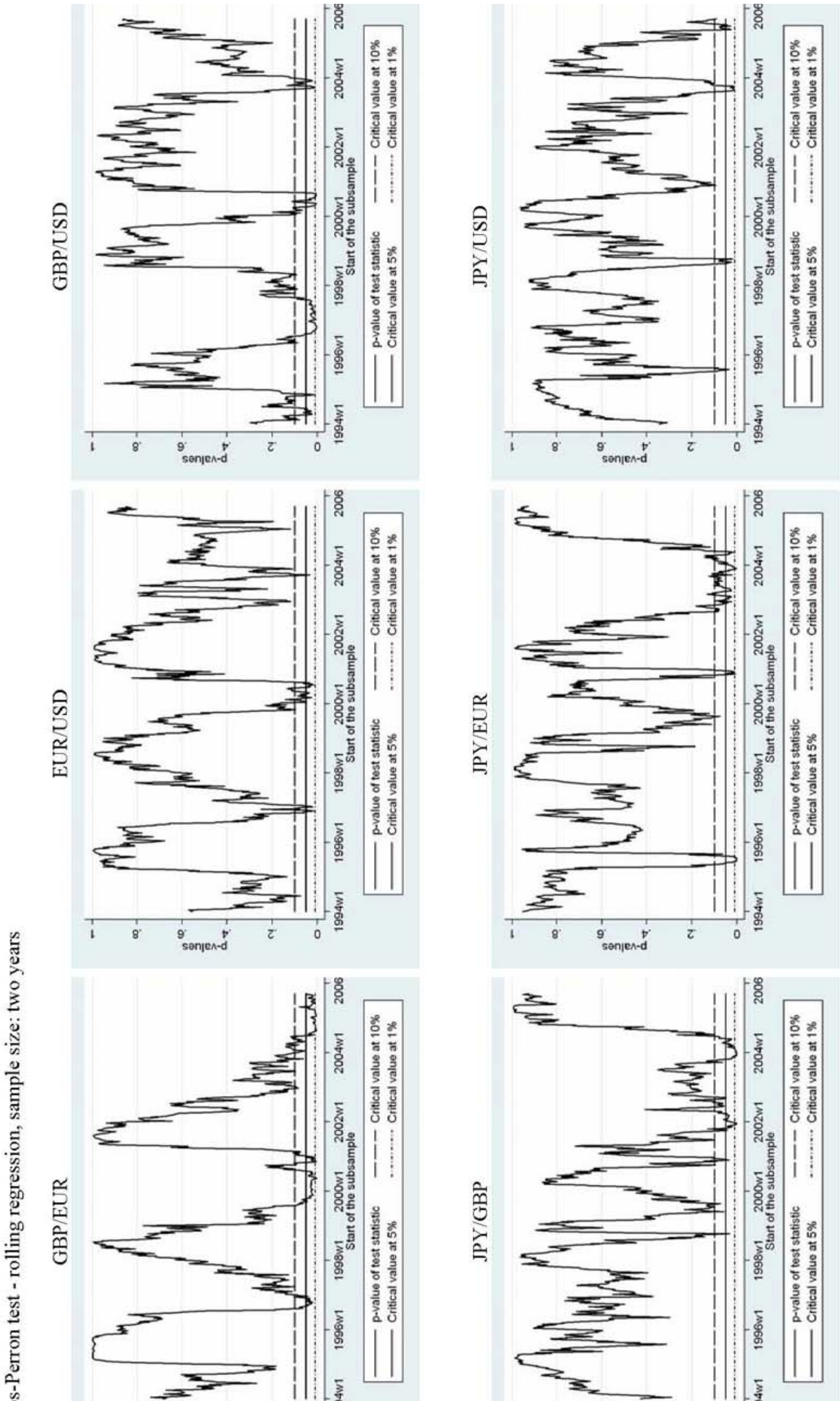

Figure 2: Phillips/ Perron test - rolling regression with a fixed sample size of two years

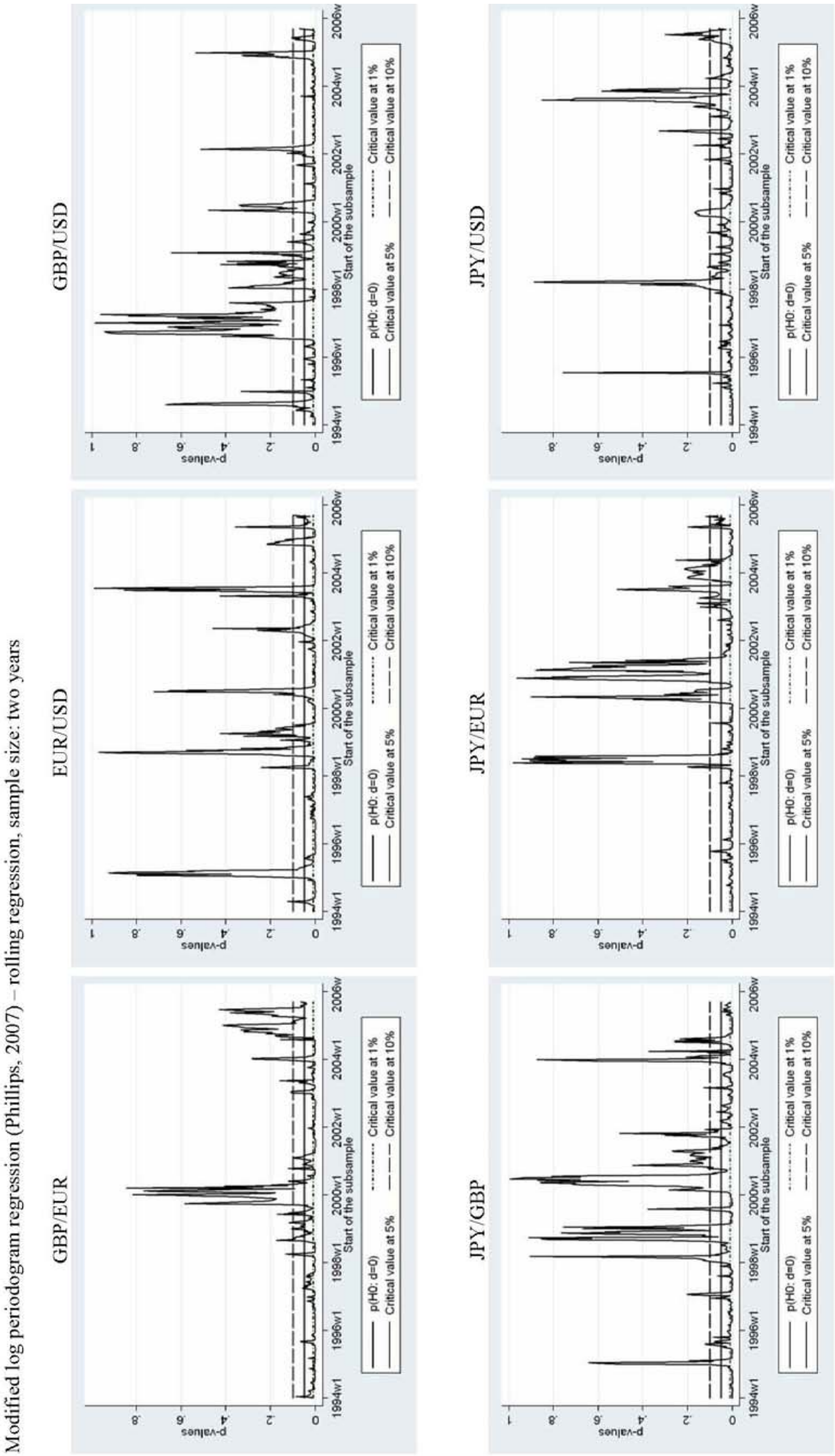

Figure 3: Modified log periodogram regression - rolling regression with a fixed sample size of two years

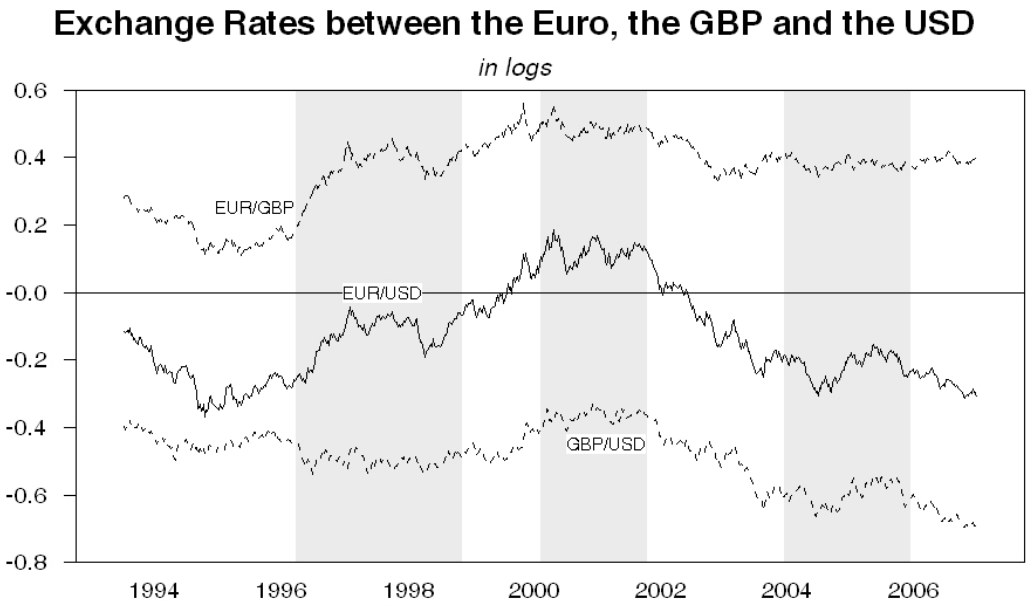

Figure 4: Exchange Rates between the Euro, the Pound Sterling and the US-Dollar

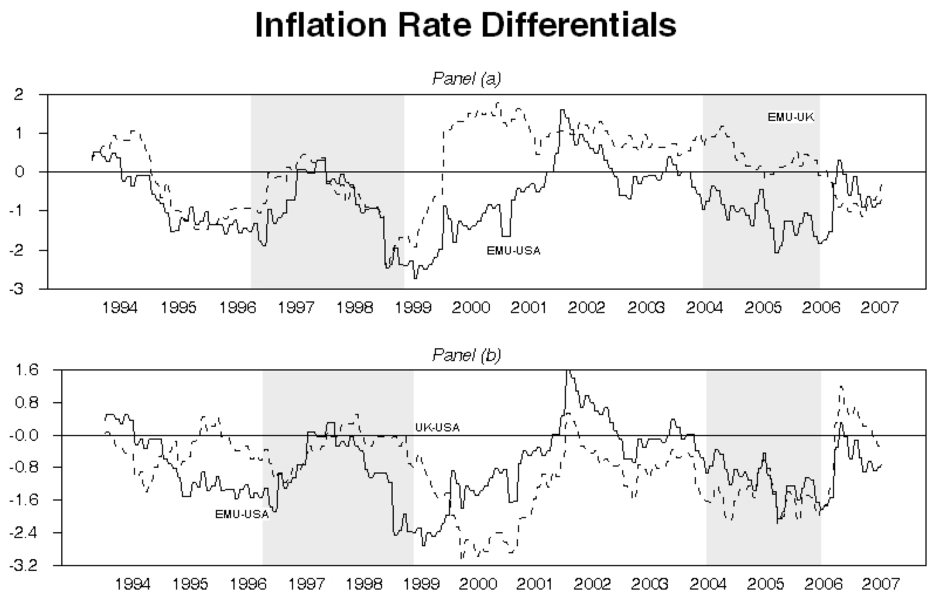

Figure 5: Differences in rates of inflation (based on monthly CPI) between the EMU and both the UK and USA (panel (a)) and between USA and both EMU and the UK (panel (b))

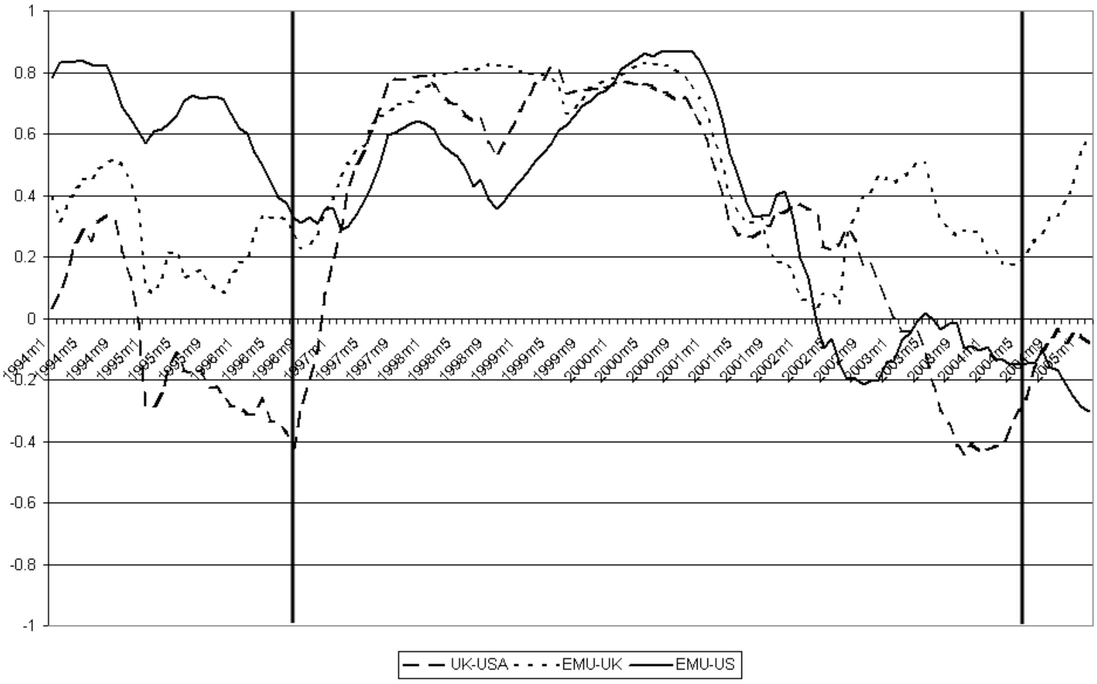

Figure 6: Correlations of the business cycle of the EMU, the UK and the USA calculated with a moving window of 3 years. The business cycle component is computed by applying the Hodrick-Prescott-filter to monthly production series that approximate real income.

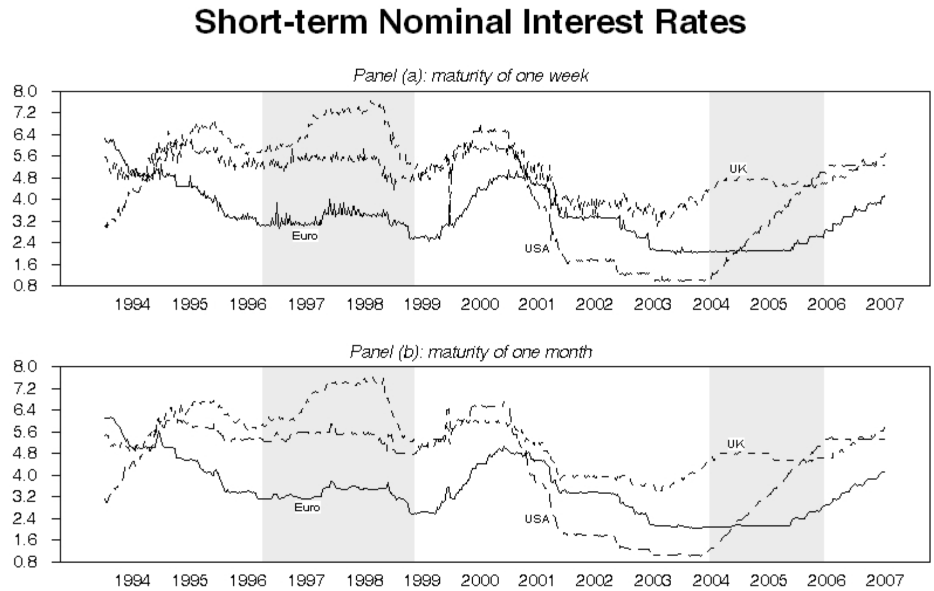

Figure 7: Short-term nominal interest rates money (market rates with a maturity of one week in panel (a) and of one month in panel (b)) of the Euro area, the UK and the USA

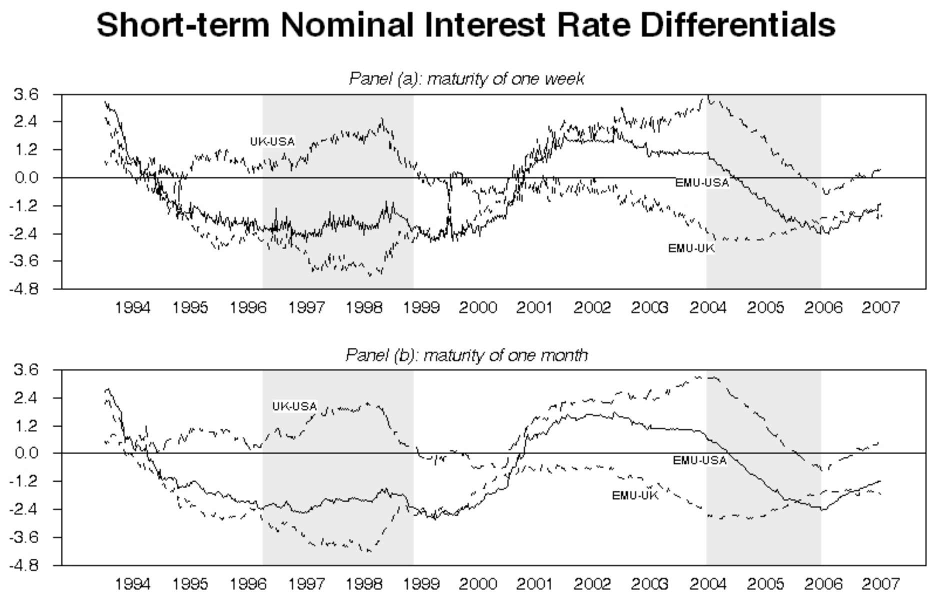

Figure 8: Differences in short-term nominal interest rates (market rates with a maturity of one week in panel (a) and of one month in panel (b)) between the Euro area, the UK and the USA