Center for European, Governance and Economic Development Research

Discussion Papers

Number 68 – October 2007

Cointegration in the Foreign Exchange Market and Market Efficiency since the Introduction of the Euro: Evidence based on bivariate Cointegration Analyses

Michael Kühl

Georg-August-Universität Göttingen

ISSN: 1439-2305

This paper is based on a presentation at the "9th Göttingen Workshop on International Economic Relations" at the Georg-August-University of Göttingen in collaboration with the Center for European, Governance and Economic Development Research (***cege***), March 22-24, 2007.

The aim of this annual workshop is to offer a forum for young researchers from the field of International Economics to present and to discuss their current topics of research with other experts. The workshop also provides the opportunity to gain an overview of recent developments, problems and methodological approaches in this field.

Detailed information on past workshops and the planning for the 2008 workshop are available at http://wwwuser.gwdg.de/~uwvw4/workshop/. Do not hesitate to contact Prof. Dr. Gerhard Rübel, ***cege*** (gruebel@uni-goettingen.de) for further questions.

Cointegration in the Foreign Exchange Market and Market Efficiency since the Introduction of the Euro: Evidence based on bivariate Cointegration Analyses

Michael Kühl*

Georg-August-Universität Göttingen

October 2007

The aim of this paper is to investigate the market efficiency on the foreign exchange market since the introduction of the Euro by applying the cointegration analysis to exchange rates. The introduction of the Euro has changed the structure of the global foreign exchange market to the extent that the second most important currency in the world with the highest credibility in the foreign exchange market, namely the Deutsche Mark, has been assimilated into the Euro. In order to evaluate if the introduction of a new currency has resulted in inefficient markets, a bivariate cointegration analysis should be applied to the seven most important exchange rates.
The empirical analysis predominantly draws on the Johansen (1988, 1991) approach and the Gregory-Hansen (1996) approach whereas the latter takes endogenous structural breaks into account. We show that the foreign exchange market is broadly consistent with the market efficiency hypothesis. A very important result is that we can find a long-run relationship between the exchange rate pairs EUR/USD and GBP/USD whereas the no-arbitrage condition is satisfied. Since the EUR/USD exchange rate is weakly exogenous the GBP/USD exchange rate takes the burden of adjustment to the long-run equilibrium.

JEL classification: C32, F31, F33, G14, G15
Keywords: Foreign Exchange Market, Market Efficiency, Cointegration

*Department of Economics, Chair in Economic Policy, Platz der Göttinger Sieben 3, D-37073 Göttingen, Fon +49(0)551/39-7335, E-Mail: michael.kuehl@wiwi.uni-goettingen.de

1 Introduction

The cointegration analysis for long-run co-movements between time series as proposed in Granger (1981, 1983) and recommended in Granger (1986) was applied to test market efficiency at the end of the 1980s. A market is said to be efficient if the price of a speculative good, like the price of an asset, reflects all available and relevant information for the pricing process. To be more precise, market efficiency is directed towards an *informationally* efficient market (Fama, 1970).
In order to test for market efficiency, the cointegration analysis is used. The argument builds on the weak form and is based on an argument highlighted by Granger (1986). On a speculative market, a pair of prices cannot be cointegrated if the market is efficient because cointegration would signify the predictability of at least one price based on the past prices of the other assets. This contradicts the weak form of market efficiency because the information set may only contain own past prices. At least two time series are said to be cointegrated if they are non-covariance-stationary[1] and integrated of the same order, and if a linear combination of them that is stationary exists (Granger, 1981, 1988). According to Granger's representation theorem, cointegration means that two or more time series fluctuate conjointly in a long-run relationship that can be seen as an equilibrium relationship, to which an error correction term automatically belongs (Granger, 1983; Engle/Granger, 1987). Short-run deviations from that long-run relationship result in an automatic adjustment process that causes the variables to return to their long-term equilibrium relationship. Thus, the error correction term contains information regarding the future movements of one variable based on past prices.
The application of the cointegration analysis has some advantages over other approaches but also investigates a different objective. Contrary to other market efficiency tests, risk premia can be neglected (Copeland, 1991).[2] Additionally, by referring to the weak form, problems concerning the joint hypothesis problem can be ruled out.[3] MacDonald and Talyor (1989), Hakkio and Rush (1989), and Baillie and Bollerslev (1989) were the first authors to apply the argumentation of Granger (1986) to the foreign exchange market in order to investigate market efficiency. The first two contributions used the Engle/Granger approach and mostly could not find any cointegration relationship.[4] MacDonald/Taylor (1989) are only able to reject the null hypothesis of no cointegration for the exchange rate pairings French Franc/ US-Dollar and Deutsche Mark/ US-Dollar. Baillie and Bollerslev (1989), however, find cointegration in a sample of seven exchange rates. Nevertheless, contrary to the two other authors, they apply the multivariate Johansen procedure. In some instances, the pappers differ significantly. Besides the application of different approaches, the frequency of the data as well as the period of observation are mostly different. While the first two contributions investigate the period after the breakdown of Bretton Woods until the mid 1980s using monthly rates; Baillie and Bollerslev (1989) only observe the first half of the 1980s using daily exchange rates. Further papers do not show any coherent tendency for different periods of observation and different approaches.
Coleman (1990), Copeland (1991), Tronzano (1992), Lajaunie and Naka (1992), Diebold et al. (1994), Lajaunie et al. (1996), and Rapp and Sharma (1999) are able to reject cointegration for

[1]Covariance-stationary is shortened to "stationary" hereafter.
[2]Although Crowder (1994) highlights the existence of a time-invariant risk premium. However, according to the literature this issue should first be neglected.
[3]A joint hypothesis problem arises for example when the market efficiency test is applied within a framework that calls for a structural model. Consequently, a joint test is carried out to test the correct structural model and market efficiency simultaneously.
[4]While MacDonald/Taylor (1989) look at more exchange rates, Hakkio/Rush (1989) only test for cointegration between the Deutsche Mark/ US-Dollar and British Pound/ US-Dollar exchange rates.

periods up to the end of the 1990s. Masih and Masih (1994) and Crowder (1994) reject the null hypothesis of no cointegration on a US-Dollar basis and Karfakis and Parikh (1994) on the basis of the Australian Dollar.[5] Some other contributions, such as Sephton and Larsen (1991), Norrbin (1996), Lajaunie and Naka (1997), and Barkoulas and Baum (1997), cannot provide a clear answer to the question of market efficiency on foreign exchange markets using the cointegration approach. As demonstrated by Sephton and Larsen (1991) and Barkoulas and Baum (1997), evidence favouring the rejection of no cointegration largely depends on the chosen period of observation.

In addition to this first strand in the literature concentrating on market efficiency, a second strand also applies the cointegration analysis to the foreign exchange market but has a different focus. Cointegration between exchange rates expressed in the same currency (that are non-stationary as a necessary condition) automatically means that the cross rates[6] have to be stationary. In a monetary system with target zones, exchange rates expressed in a currency that does not belong to it shall be cointegrated. The cross rates should therefore move within the defined ranges and exhibit mean-reverting behaviour. Consequently, a cointegration analysis can be taken to evaluate the stability within a monetary exchange rate system. By applying a multivariate framework, the exchange rates should share common stochastic trends whereas the number declines if a high integration of the system is achieved. With respect to the Exchange Rate Mechanism (ERM) in particular, a couple of authors tested for stability within the European Monetary System (EMS). Both Hakkio and Rush (1989) and Copeland (1991) have already stressed that comovements can be expected between fixed or managed floating exchange rates. Norrbin (1996), Woo (1999), Haug et al. (2000), Rangvid and Sorensen (2002), and Aroskar et al. (2004) are mostly able to reject the null hypothesis of no cointegration for the EMS currencies. In particular, strong evidence supporting the rejection of no cointegration is present in the period before the introduction of the Euro, where parities were strongly fixed. In contrast, the evidence in favour of no cointegration was less pronounced in the mid 1980s. Taking the realignments during that time into account, the evidence increases dramatically, as illustrated by Woo (1999). The introduction of potential structural breaks especially improves the evidence against no cointegration. Jeon and Seo (2003) and Phengpis (2006) examine structural instability in particular and investigate currency crises for market efficiency. The results are mixed. No paper has explicitly examined the foreign exchange market for cointegration since the establishment of the European Monetary Union. Hence, the aim of this paper is to investigate the market efficiency of the foreign exchange market since the introduction of the Euro using the cointegration methodology. The introduction of the Euro has changed the structure of the global foreign exchange market to the extent that the second most important currency in the world with the biggest credibility on the foreign exchange market, namely the Deutsche Mark, has been assimilated into the Euro (BIS, 2005). In order to evaluate whether the introduction of a new currency that has commonly replaced more established and less established currencies has resulted in inefficient markets, the cointegration analysis should be applied to the seven most important exchange rates: the Australian Dollar, the Canadian Dollar, the Swiss Franc, the British Pound Sterling, the Euro, the Japanese Yen, and the Swedish Krona. In keeping with the literature, all currencies are expressed in US-Dollar. Contrary to the literature, where primarily a set of foreign exchange rates is tested for cointegration simultaneously, the cointegration analysis is carried out on a bivariate basis. The reason for which comovements between two exchange rates should be examined explicitly is that Granger's argumentation in his original paper of 1986 is applied to two prices on a speculative

[5]Contrary to other works Karfakis and Parikh (1994) take the AUD as the base currency.
[6]The division of the first one by the second.

market. Information on the future developments of one price is only taken from the second price.
The remainder is organised as follows: after a short introduction into the methodology of cointegration and the Johansen approach in section two, the link between cointegration and market efficiency is discussed in the third section. In doing so, a caveat concerning the role of common fundamentals is pointed out. Since the literature does not discuss the choice of the model underlying the estimation in more detail, this shall be done more extensively in section 4. The fifth section presents the empirical results of the Johansen approach and a second cointegration technique taking structural breaks into accounts, namely the Gregory-Hansen-approach.

2 Econometric Methodology

2.1 Definition of Cointegration and the Engle/Granger approach

If time series have the same order of integration and if a linear combination of these time series exist that is stationary (integrated of order one), these series are referred to as being cointegrated. Originally, Engle and Granger (1987) proposed a two-step procedure to estimate cointegration relations. In this two-step procedure, which is mostly applied to only two time series, the first series (y_t) is regressed on the second series (x_t). After the first step, the resulting error series (z_t) is tested for stationarity (see equ. (1)).

$$y_t = \mu + bx_t + z_t \tag{1}$$

If the null hypothesis of non-stationarity is rejected, it can be said that the time series are cointegrated and b is the cointegration parameter (Engle/Granger, 1987).[7] As a result, the null hypothesis is equivalent to the statement of no cointegration.
Cointegration means that both series move together in the long-run and cannot drift apart very much from each other (Granger, 1981). Consequently, the error term resulting from the linear combination of the aforementioned time series can be seen as an equilibrium error. This equilibrium error quantifies the deviation of the time series from their common long-run relationship (Granger, 1986). Hence, deviations from the linear combination can only occur randomly and unsystematically.
If a long-run relation between time series processes exists and if the equilibrium error is stationary, there must be a mechanism that brings the system back to equilibrium in the face of an innovation. In this context, Granger (1983) analytically illustrates that the consideration of a cointegration relationship is equivalent to the existence of an error correction mechanism (Granger's representation theorem).

2.2 Johansen Approach

In the one-equation approach, a problem arises due to the formulation of the regression equation. Using the Engle/Granger procedure, the choice of the independent and the dependent variable must be made previously. Basically, both results have to coincide in the statistical limit. However, working typically with finite samples, the correct choice of the dependent variable is important. In order to avoid misleading results in finite samples, the procedure has to be carried out twice, with a different formulation in the second run. If more than two series are under observation, the computational efforts increase considerably. In addition, the previous determination of the

[7] The coefficient μ is an intercept.

independent and dependent variables excludes possible endogeneity among the variables.
A different method to test for cointegration is based on a vector autoregressive (VAR) model. The vector X_t contains the endogenously seen variables and has the dimension $n \times 1$, where n is the number of endogenous variables. Each variable follows a process that is influenced by its own lagged variables and the lagged variables of the other endogenous variables.

$$X_t = \Pi_1 X_{t-1} + ... + \Pi_k X_{t-k} + \epsilon_t \quad \text{with} \quad t = 1, ..., T. \tag{2}$$

The matrix of coefficients Π_k has the dimension $n \times n$. Based on (2), the VAR can be transferred to a VAR of first differences. For this purpose, the lagged variables of the endogenous variables are subtracted from both sides and the following system arises

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \epsilon_t, \tag{3}$$

whereas $\Gamma_i = -I + \Pi_1 + ... + \Pi_i$ with $i = 1, ..., k-1$ and $\Pi = -(I - \Pi_1 - ... - \Pi_k)$ (Johansen/Juselius, 1990, p. 170.) Here the matrices Γ_i contain information on the short-run adjustment coefficients of the lagged differenced variables. Additionally, the expression ΠX_{t-1} indicates the error correction term, i.e. it includes the long-run relationships between the time series (c.f. Lütkepohl, 2005, pp. 247-249; Johansen, 1992b, p. 315). Using the matrix Π, further conclusions regarding the number of cointegration relations can be made if the rank of the matrix is known (Johansen/Juselius, 1990, p. 170).[8]
The Johansen procedure adopts the idea of determining the rank of the matrix Π. In general, the rank of a matrix shows the number of linearly independent processes that is equivalent to the number of linearly independent columns. According to the definition, departing from the relevant case of I(1) variables in levels, both the differences of the endogenous variables and their lagged differences are stationary.[9] For this reason, a test for cointegration aims at testing the rank of Π. If the rank of the matrix Π is greater than zero and less than the number of endogenous variables n, the matrix with the dimension $p \times r$ can be decomposed into the matrices α and β, so that $\Pi = \alpha\beta'$. Using the cointegration vector β, the non-stationary vectorprocess X_t can be made stationary by generating linear combinations $\beta' X_t$ (Johansen, 1988, p. 232). In this case, the system in (3) becomes a vector error correction model and, in doing so, the matrix α describes the adjustment speed for each variable after a deviation from the long-run relationship. In other words, the elements in α weight the error correction term in each row of the VECM. Furthermore, the matrix β contains the coefficient of the cointegration relation, i.e. the weights within the linear combination. Subsequently, the VECM is a reduced form of the VAR in (2). Only the hypothesis of a restricted matrix Π is implemented. The cointegration rank can be tested by using the procedures outlined by Johansen (1988, 1991). On the basis of these considerations, the test statistics for the statistical significance of the rank of the matrix Π can be derived (Johansen, 1995, p. 89-95).
The first test weights the hypotheses of, at most, r cointegration vectors, i.e. $Rank(\Pi) = r$, against the alternative of $Rank(\Pi) > r$, that is to say, there are r or more cointegration relations. According to Johansen (1988, 1991) this test is based on a likelihood ratio test and is called *"'trace statistic"'*.

$$\lambda_{trace} = -T \sum_{i=r+1}^{p} \ln(1 - \hat{\lambda}_i). \tag{4}$$

[8]The rank of the matrix is equivalent to the number of cointegration vectors.
[9]The error term is also stationary by definition.

Additionally, Johansen proposes a second test to determine the cointegration rank. As the first test, it is also based upon a likelihood ratio test but can differentiate more precisely between two alternatives, i.e. the ranks of the matrix Π. This means, it is tested if there are exactly r cointegration relations or if there is just one more. Since this test departs from the eigenvalues that are arranged by their magnitude, the test is called *"'maximum eigenvalue test"'*.

$$\lambda_{max} = -T \ln(1 - \hat{\lambda}_{r+1}). \tag{5}$$

Both test statistics are distributed asymptotically as χ^2 with $p - r$ degrees of freedom (Johansen/ Juselius, 1990, pp. 177-179; Johansen, 1991, pp. 1555/1556). As suggested by Johansen and Juselius (1990), both test statistics should be used simultaneously, although different conclusions can be drawn. In order to estimate the parameters like the cointegration vector, adjustment coefficients or eigenvalues, the Maximum Likelihood Procedure is applied.

3 Marketefficiency and Cointegration

3.1 The Market Efficiency Hypothesis

A widely cited definition of market efficiency traces back to Fama's survey article in 1970. According to him, a market is deemed efficient if the prices on that market fully reflect all available information relevant for the pricing process (Fama, 1970, pp. 383/384). More precisely, the expression "efficient market" refers to an *informationally* efficient market. As Jensen (1978) writes, a market is - related to a specific information set (Φ_t) - efficient if none of the market players can earn excess profits by exploiting the known information set (Jensen, 1978, p. 96). Fama (1970) subdivides into three categories, namely the *weak form*, the *semi-strong form* and the *strong form*. In the weak form, the information set only comprises past prices. Consequently, the information set contains all information that is included in historical prices. In the semi-strong form, the information set additionally comprises all publicly available information relevant for the pricing process. In particular, the fundamentals determining the price belong to this category. Finally, the information set in the strong form also includes private information. Thus, a market is said to be efficient if trading on the basis of private information cannot yield higher profits.

In addition to the weak form, the semi-strong form includes information on the fundamentals forming the price. Departing from the semi-strong form a full, specified market model is necessary to evaluate the correct impact of the fundamentals and, hence, the correct price formation. Thus, a test for market efficiency automatically tests the correct market model as well. As a result, in this manner a test for market efficiency always tests a joint hypothesis implicitly. Inferences regarding the market efficiency cannot automatically be drawn from the rejection of the hypothesis because a rejection can be due to an unspecified market model and not due to market inefficiencies (Fama, 1970, pp. 1022-1025; Fama, 1976, pp. 136/137).

3.2 The Coherence between EMH and Cointegration

Basically, the application of the weak form avoids the problem of joint hypotheses. In this fashion, a test for market efficiency that does not require the specific formulation of an equilibrium price mechanism goes back to an argument by Granger (1986) and aims at the development of two or more (asset) prices. If two or more asset prices show a stable common relationship in the long-run, i.e. if two or more asset prices are cointegrated, it is possible that the movement of one asset price is linked to the movement of other asset prices. As already described above, the establishment of a cointegration relationship is equivalent to the existence of an error correction term. In this case, the price of one asset does not only depend on its own past prices but also on the history of a different asset's prices. Thus, the weak form of market efficiency is violated (Richards, 1995, p. 632). The error correction term implies that in the face of a deviation of one asset price from the induced long-run relationship, unused profit opportunities would automatically arise. If the stable long-run relationship between prices is known to the market participants they are able to exploit them and are in position to make excess profits (Copeland, 1991, p. 187).

MacDonald and Taylor (1989) were amongst the first to apply the cointegration methodology to testing for market efficiency on the foreign exchange market. A small formulation of the interrelation between market efficiency and cointegration is provided at the beginning of their contribution. This will be adapted in the following. A three-country case is assumed where the exchange rates s_t^{ij} are expressed in the same currency. The exchange rate is the domestic currency i in terms of the foreign currency j. If cointegration can be established, it follows that:

$$\Delta X_t = \alpha(\beta' X_{t-1}) + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \epsilon_t, \tag{6}$$

with $X_t = [s_t^{12}, s_t^{13}]'$ and $\beta' = [1, -b]$ so that the error correction term (ect) is $s_t^{12} - b \cdot s_t^{13}$. In equ. (6) the same error correction term $(s_{t-1}^{12} - b \cdot s_{t-1}^{13})$[10] is included in both equations and thus determines both exchange rates. Here, the error correction term shows the long-run dynamic of the two exchange rates. In the α-vector, the coefficients describing the adjustment process due to a disequilibrium can be seen in particular.[11]

The existence of cointegration between exchange rates contradicts the weak form of market efficiency because of its forecastability. We can assume that the expectations regarding the exchange rate s_t^{12} in period t are based upon the information set Φ_{t-1}^{12}.[12] After the inclusion of the past realisations of the exchange rates s_t^{13}, the information set should not be expanded qualitatively if the market is efficient. Taking the information set Φ_{t-1}^{12} and adding the information set Φ_{t-1}^{13}, the exchange rate s_t^{12} no longer depends solely on its own past realisations but also on the historical realisations of the exchange rate s_t^{13}.

Consequently, the information set of s_t^{12} was formally extended quantitatively and composes both information sets:

$$\Phi_{t-1} = \Phi_{t-1}^{12} + \Phi_{t-1}^{13}. \tag{7}$$

If a market is efficient, the expectations regarding the exchange rate s_t^{12} must not differ with respect to the two information sets. It follows in the style of MacDonald/Taylor (1989)

$$E(s_t^{12} | \Phi_{t-1}^{12}) = E(s_t^{12} | \Phi_{t-1}) \tag{8}$$

with

$$\Phi_{t-1}^{12} = \{s_{t-1}^{12}, s_{t-2}^{12}, s_{t-3}^{12}, ...\} \tag{9}$$

and

$$\Phi_{t-1} = \{s_{t-1}^{12}, s_{t-2}^{12}, s_{t-3}^{12}, ..., s_{t-1}^{13}, s_{t-2}^{13}, s_{t-3}^{13}, ...\} \tag{10}$$

[10] b is the cointegration parameter.
[11] Whereas $\alpha_1 < 0$ and $\alpha_2 > 0$
[12] Φ_{t-1}^{ij} will only contain past prices, i.e. past exchange rates s_t^{ij}.

If the error correction term in equ. (6) is valid equ. (8), i.e. the EMH, can no longer be held. If cointegration can be observed, the quantitative expansion of the information set coincides with the qualitative expansion of the information set.
Using the error correction term, one exchange rate can be predicted by using the other if the long-run relationship and past exchange rates are known to the market. This means that causality runs at least in one direction. When it comes to the weak form the market is not efficient.[13]

3.3 Cointegration and Variations in Common Fundamentals

As widely and controversially discussed in the literature, the exchange rate can be explained by different theories.[14] Depending on the particular theory, different fundamentals help explain the exchange rate. Subsequently, not only are past prices important in evaluating the market efficiency; the fundamentals also play a significant role. This means a departure from the weak form of EMH. Due to the problems that arise using the semi-strong form concerning the joint hypothesis problem, the application of the semi-strong form in the face of fundamentals is inadequate. However, the impact of fundamentals on the eyed exchange rates cannot be excluded in an empirical investigation. Exchange rates are nothing more than relative prices. Contrary to asset prices taken from the capital market, exchange rates are determined by fundamentals from two different destinations. Treating the three-currency (country) case again, three exchange rates (s_t^{12}, s_t^{13} and s_t^{23}) exist which are assumed to be flexible. Based on these considerations, each subset of fundamentals only contains the fundamentals of one country Φ_t^i. Summing up the individual information sets yields the common information set Φ_t.

$$\Phi_t = \Phi_t^1 + \Phi_t^2 + \Phi_t^3 \tag{11}$$

Each expected exchange rate is based upon its fundamentals included in the corresponding information set known at time $t-1$.

$$E(s_t^{12}|\Phi_{t-1}^1, \Phi_{t-1}^2) = s_t^{12} \tag{12}$$

$$E(s_t^{13}|\Phi_{t-1}^1, \Phi_{t-1}^3) = s_t^{13} \tag{13}$$

$$E(s_t^{23}|\Phi_{t-1}^2, \Phi_{t-1}^3) = s_t^{23} \tag{14}$$

Only the cases (12) and (13) are relevant for further explanations. As can been seen very easily, the expectations regarding the exchange rate for both exchange rates are built on the fundamentals of country 1, i.e. the information set comprising the fundamentals of country 1. If the fundamentals of country 2 and 3 remain unchanged and assuming that an innovation in the fundamentals of country 1 takes place, it is straightforward that the exchange rates s_t^{12} and s_t^{12} move in the same direction. Hence, the exchange rates are influenced by variations in common fundamentals.[15] If changes in common fundamentals predominate, movements in the same direction should occur. Baffes (1994) and Ferré and Hall (2002) show this in a similar manner. An investigation of the foreign exchange market for the purpose of testing market efficiency applying the cointegration analysis is inadequate to the extent that in an ex post treatment the empirical analysis cannot differentiate between the rejection of the weak form and the importance of variations in common fundamentals. If no cointegration between exchange rates can be rejected, the consideration of

[13] The short-run dynamic via lagged differenced variables in the VECM is neglected.
[14] For a survey see Taylor (1995).
[15] Crowder (1994) and Barkoulas et al. (2003) argue that a variation in the risk premium results in the consideration of cointegration. This argument can be subsumed under this proposition.

variations in common fundamentals is inevitable. Only if variations in common fundamentals can be excluded as driving forces behind common movements is cointegration incompatible with market efficiency.

Hakkio and Rush (1989) have already highlighted this issue. They argue that exchange rates cannot be cointegrated if they are "different assets". Copeland (1991) refers to a similar point.[16] Another source of cointegration stems from a system of fixed exchange rate regimes or from a target zone regime.[17]

In the case of a target zone regime, bands for the exchange rates are defined. The central banks of participating countries have to intervene if the exchange rate drops out of the committed ranges. Consequently, within this range the exchange rates are more or less flexible, but when achieving the upper or lower band, interventions cause the rates to remain within the band. If these particular currencies are expressed in the same currency not belonging to the system, cointegration between these exchange rates occurs due to the interventions of the central banks. Norrbin (1996), for example, picks up this point and abandons the efficient market argument but also applies the cointegration methodology to exchange rates.[18] Cointegration would mean that the cross rates, i.e. the exchange rates belonging to a monetary system, should not fluctuate too heavily, i.e. they should be stationary. Despite the different orientation of that particular investigation, again common fundamentals, i.e. interventions, are the driving forces.[19]

Despite the problems stemming from the use of cointegration analysis on foreign exchange markets due to variations in common fundamentals, the methodology does not have to be changed. As Frenkel and Levich (1975) and Levich (1985) suggest, if transaction costs are neglected, a specific amount of money in currency 1 has to retain its value, even if it is converted across the two other currencies. They refer to this as *triangular arbitrage*. Dwyer and Wallace (1992) pursue the same argument and demonstrate that the foreign exchange market in a three-country case is efficient if no cross-sectional arbitrage opportunities exist.[20] This implies that the proportion of two exchange rates based on the same currency has to equal the cross rate. Turning to exchange rates in logarithms market efficiency means

$$s_t^{12} - s_t^{13} = s_t^{32}. \tag{15}$$

Equ. (15) describes the so-called *no arbitrage* condition without transaction costs. A market is efficient despite the fact that cointegration cannot be rejected if the cointegration vector is $\beta' = (1, -1)$. For this purpose the exchange rates s_t^{12} and s_t^{13} have to be integrated of order one.[21] In addition, it follows that the cross rate is integrated of order zero (Dwyer/Wallace (1992), p. 321).[22]

4 Matching up the theoretical foundation and the Johansen approach

Johansen proposed the VECM as it is outlined in equ. (3) in his article of 1988. Baillie and Bollerslev (1989) were the first authors to examine the foreign exchange market for cointegration

[16] In this respect, if variations in common fundamentals are responsible for a long-run relationship two exchange rates are not different assets.

[17] Engel (1996) illustrates this with a monetary rule, explicitly taking a target for the exchange rate into account.

[18] Haug et al. (2002) investigate the number of common stochastic trends. In a bivariate setting this investigation is equivalent with the outlined argument.

[19] Similarly, internationally organised interventions, such as the Plaza Agreement and Louvre Accord can be the reasons.

[20] Rapp/Sharma (1999) speak about efficiency *across* countries when testing for co-movements among exchange rates.

[21] This test can also be seen as a test for the same information set regarding fundamentals. In this respect, the application of the cointegration methodology implicitly tests for different asset where the null hypothesis is equivalent with the null hypothesis of no cointegration.

[22] This is a special case of the general solution where all three exchange rates are seen as non-stationary as outlined by Baffes (1994). See also Ferré/Hall (2002), p. 134.

using the Johansen approach and refer to Johansen (1988). An extension and generalisation of the procedure was developed by Johansen and Juselius (1990) and Johansen (1991). In the updated setting, deterministic components were introduced.[23]

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \mu + \Phi D_t + \epsilon_t. \tag{16}$$

with

$$\Pi X_{t-1} = \alpha(\beta' X_{t-1}) \quad \text{and} \quad \mu = \mu_0 + \mu_1 t \tag{17}$$

where $\mu_1 t$ is a linear trend in differences.

Johansen (1994) provides a detailed discussion on the role of the constant term. He demonstrates that the VECM in equ. (16) can be rewritten using Granger's Representation Theorem with the representation

$$X_t = C \sum_{i=1}^{t} \epsilon_i + \frac{1}{2}\tau_2 t^2 + \tau_1 t + \tau_0 + C(L)\epsilon_t + A \tag{18}$$

where $C = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}\alpha'_\perp$ with $\Gamma = I - \sum_{i=1}^{k-1} \Gamma_i$ as already defined, $\beta' A = 0$, $\tau_1 = C\mu_0$ and $\tau_2 = C\mu_1$.[24] The expression $C(L)\epsilon_t$ describes the moving average part and $\alpha_\perp$ and $\beta_\perp$ are $p \times (p-r)$ matrices of full rank. Both are orthogonal to its corresponding counterparts, the matrix of loading coefficients α and the cointegration vector β. Furthermore, the constant terms in equ. (16) and equ. (17) respectively can be decomposed into their components

$$\mu_i = \alpha\beta_i + \alpha_\perp \gamma_i \quad \text{with} \quad i = 0, 1. \tag{19}$$

β_i denotes that the expression belongs to the cointegration space and results as $\beta_i = (\alpha'\alpha)^{-1}\alpha'\mu_i$. γ_i describes the part of the deterministic components that are outside the cointegration space and is $\gamma_i = (\alpha'_\perp \alpha_\perp)^{-1}\alpha'_\perp \mu_i$. For reasons of simplicity and due to the purpose of this section, τ_2 is restricted to zero, i.e. quadratic deterministic trends are excluded.[25] Consequently, equ. (19) is restricted to the case $i = 0$. Hence, $\alpha'_\perp \mu_1$ is zero or equivalently $\gamma_1 = 0$. Three different cases can therefore be considered.

$$\begin{aligned} H_1(r): &\quad \mu_t = \alpha\beta_0 \\ H_2(r): &\quad \mu_t = 0 \\ H_3(r): &\quad \mu_t = \alpha\beta_0 + \alpha_\perp \gamma_0 \end{aligned} \tag{20}$$

Now, two different cases with respect to the deterministic component have to be distinguished in addition to the case of no constant term (H_2). If τ_1 is unequal to zero, it is equal to $C\mu_0 = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}\alpha'_\perp \alpha_\perp \gamma_0$. It can be seen that a linear trend is present and cancels out in the cointegration relationship but is still present in the VECM (H_3) in the form of a drift term since $\gamma_0 = (\alpha'_\perp \alpha_\perp)^{-1}\alpha'_\perp \mu_0$. Additionally, restricting $\alpha'_\perp \mu_0$ to zero, the linear trend vanishes but $\alpha\beta_0$ is still present and the cointegration space consists of a constant term (H_0) (Johansen, 1994, pp. 206-210; Johansen, 1995, pp. 80-84).

Thus, Johansen illustrates two different meanings with respect to the constant. In the first, the constant term represents a linear trend in levels. If the constant in differences does not generate

[23] ΦD_t represents further deterministic components such as seasonal dummy variables.

[24] A is the initial value.

[25] The possibility of a quadratic linear trend in levels that is common to both exchange rates in the analysis is seen as negligible.

a linear trend in levels, it simply represents a non-zero mean and becomes an intercept in first differences, as well as in levels, and can be absorbed into the cointegration space.
While Baillie and Bollerslev (1989) are able to reject the null hypothesis of no cointegration for a multivariate set of seven exchange rates, Diebold et al. (1994) do not confirm their results by using the same sample period and applying the Johansen (1991) setting, i.e. highlighting the importance of the constant term. A similar result with different sample periods is obtained by Barkoulas and Baum (1997), again by taking the constant into account.[26] We should bear in mind the fact that the inclusion of a constant term in the cointegration approach tends to result in a non rejection of the null hypothesis of no cointegration.
Unfortunately, the precise specification of the underlying regression model is mostly not attempted. Barkoulas and Baum (1997) speak about a drift term in the VAR process and include a constant in the VECM (pp. 636) representing a linear trend in levels. As far as the constant term is concerned, Diebold et al. (1994) argue that a constant should be borne in mind as long as there is no reason to exclude it (p. 6). MacDonald and Marsh (1999) also explicitly introduce constant terms in the underlying VAR. In the EMH cointegration literature, only Lajaunie and Naka (1992), Lajaunie et al. (1996) and Lajaunie and Naka (1997) lay more emphasis on the role of the constant and test for linear trends.
All the aforementioned works have the multivariate analysis in common. Norrbin (1996) investigates bivariate cointegration among exchange rates within the EMS and also explicitly includes a constant term. According to him, each "... series may also contain a deterministic drift" (p. 1507). However, his investigation is directed towards the stability of the EMS. The reason to include a drift stems from the observation that the EMS rates drift apart deterministically. Therefore, a constant term is integrated in the VECM formulation that also represents a trend in levels.
Unfortunately, a more detailed investigation of the constant term in the face of EMH considerations is only provided by few of the authors. Departing from a linear trend in levels, a non stochastic process drives the system of exchange rates. According to the EMH, the deterministic trend should be observed by the market participants. Without cointegration the market is efficient if deviations from the linear trend are not predictable. The linear trend itself does not cause the market to be inefficient because it is included in the information sets of both exchange rates so that the same linear trend can be observed. Using the linear trend, the future exchange rates can be predicted simultaneously. This should contradict the EMH with respect to one asset price by using the paradigm of the pure random walk model[27] but not by applying the martingale model.[28] Hence, market efficiency is not affected in the sense meant by Granger (1986) because the common linear trend can be seen as a fundamental variable that is common to both exchange rates. Consequently, a deterministic trend does not contradict market efficiency as long as cointegration is not present and all market participants have recognised the linear trend.[29]
The period of observation in Baillie and Bollerslev (1989) and thus in Diebold et al. (1994) comprises the first half of the 1980s. During this period, the US-Dollar appreciated strongly against

[26] Baillie/Bollerslev (1994) reinforce the consideration of no cointegration and apply the methodology of fractional cointegration (see for fractional integration Granger/Joyeux (1980)). Baillie/Bollerslev (1994) find that the deviations from the long-run relationship follow a long memory process ($I(d)$ and $0 < d < 1$). Granger/Hyung (2004) argue that long memory properties can result from occasional breaks. For this reason, a cointegration test taking a structural break into account is applied in section 5.4.
[27] Further details can be found in Fama (1965a,b, 1970). A statistical equivalent could be a random walk with drift at the most. Here, the time series would wander randomly and not predictably around the linear trend.
[28] Arguing in line with LeRoy (1985), a linear trend would automatically not contradict EMH because it is possible that the equilibrium expected price is not constant. See also Levich (1979, 1985).
[29] This is a necessary condition for market efficiency. The linear trend should belong to the information that is publicly available.

other major currencies. The comparison of the results of Baillie and Bollerslev (1989) and Diebold et al. (1994) shows that common movements seem to prevail. Instead of cointegration, a linear trend tends to be responsible for that finding. If the constant in the VECM does not lie within the cointegration space, it generates a linear trend in levels that is common for all observed variables and hence not relevant for the long-run relationship (Juselius, 2006, p. 95-100). Lajaunie and Naka (1992) examine the second half of the 1980s, when the US-Dollar depreciated strongly against other major currencies due to the joint interventions of the central banks in industrialised countries. For this period, they cannot find cointegration in a multivariate setting either. Additionally, they are able to reject the absence of a linear trend, which is also confirmed by Lajaunie et al. (1996).[30] However, Haug et al. (2000) test explicitly for the correct specification on the basis of the proposed models in Johansen and Juselius (1990) and Johansen (1991). With their included exchange rates and period of observation, a model without a deterministic term is the optimum in a multivariate setting (Haug et al., 2000, p. 426).[31]

It should be reemphasized that the VAR including an intercept is equivalent to the formulation of the VECM with an intercept if the constant term in first differences does not generate a linear trend in levels (Lütkepohl, 2005, p. 257). In this case, the constant term can be absorbed into the error correction term. As a result, there is a difference between the case where a constant term in the VECM represents a linear trend in levels and the case where it represents a non-zero mean both in first differences and levels. Statistically, two models arise with different critical values and different test statistics (Johansen/Juselius, 1990; Johansen, 1991; Osterwald-Lenum, 1992).

Economically, the formulation of the model has important implications for testing the market efficiency hypothesis, especially in the bivariate setting. Here, a differentiation must be made between the case in which the intercept of the cointegration relationship is zero and the case in which it is different from zero. The outlined no-arbitrage condition in eq. (15) does not take a constant term into account because a constant term would have the interpretation of transaction costs. However, a constant term cannot be neglected in the empirical analysis that starts from eq. (15) and (16). If two exchange rates are cointegrated in an efficient market environment, i.e. the cointegration vector is $(1, -1)$, the left hand side of eq. (15) will be the error correction term as long as the values of s_t^{12} and s_t^{13} coincide at each point in time. If they differ, it will be necessary to introduce a constant term that reflects the difference in their levels.[32] Hence, the constant term is necessary to generate s_t^{12} from s_t^{13} or vice versa in that case. Consequently, a constant term that represents a non-zero mean has to enter the cointegration equation, i.e. the error correction term. Thus, model 1 is to apply instead of model 2. In the formulation of model 1, the cointegration relationship consists of

$$ect = \mu_0 + s_t^{12} - s_t^{13}. \tag{21}$$

In this formulation, μ_0 is assumed to be deterministic, i.e. the non-zero mean does not change over time. The cointegration vector β consistent with EMH does not change but the error correction term is amended by a constant term, $\alpha(\beta' X_{t-1} + \mu_0)$ with $\beta' = (1, -1)$. The no-arbitrage condition still remains the same.

Furthermore, a time trend in levels cannot be excluded a priori. Situations with a time trend in levels can coincide with market efficiency. As long as no answer is given to the open questions concerning the intercept and the time trend in levels, all three cases must be taken into account:

[30] Concerning the linear trend, they argue that the linear trend can be seen as a "market drift which may not be predictable", Lajaunie et al. (1996), pp. 557-561.

[31] As already mentioned, Haug et al. (2000) investigate a different issue and thus use the Deutsche Mark and ECU respectively as the base currency.

[32] In this respect, the constant term represents a non-zero axis intercept.

non-zero mean in the error correction term, zero mean, and time trend in levels. Basically, the cointegration vector must assume the form described in equ. (15) if the market, despite cointegration, is efficient. If a time trend is present and if cointegration cannot be rejected, the coefficient must take the aforementioned forms. In order to obtain the correct model, statistical inferences must be made carefully first. The following three relevant models are based on Johansen (1991, 1992a, 1994) and bear in mind the critique of Diebold et al. (1994).[33]

Model 1: non-zero mean in ect

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \epsilon_t. \tag{22}$$

with $\Pi X_{t-1} = \alpha(\beta' X_{t-1} + \mu_0)$.
Model 1 represents equ. (15) with a non-zero mean, i.e. an intercept in the cointegration relationship. The constant term is assimilated into the error correction term.

Model 2: zero mean

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \epsilon_t. \tag{23}$$

with $\Pi X_{t-1} = \alpha\beta' X_{t-1}$.
Model 2 is the original model proposed by Johansen (1988) without any deterministic components.

Model 3: linear trend in levels

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \mu + \epsilon_t. \tag{24}$$

with $\Pi X_{t-1} = \alpha\beta' X_{t-1}$.
Finally, model 3 takes a linear trend in levels into account.

5 Empirical Results

5.1 Descriptive statistics

The period under observation runs from 4 January 1999 to 29 December 2006 and covers the daily exchange rates of the US-Dollar expressed in foreign currencies. Overall, the seven most important exchange rates in the world are used: the Australian Dollar (AUD), the Canadian Dollar (CAD), the Swiss Franc (CHF), the British Pound Sterling (GBP), the Euro (EUR), the Japanese Yen (JPY), and the Swedish Krona (SEK).[34] All daily exchange rates are converted by taking the natural logarithm. The data are taken from the database of the Federal Reserve Bank of St. Louis and are noon buying rates in New York City.
Figure 1 depicts the daily exchange rates for the USD (USD in foreign currencies) normalised to the corresponding exchange rate on 4 January 1999. As can be seen, there is a parallel movement between some of the exchange rates. A narrower trend can be observed particularly since the beginning of 2002. The CHF, GBP and EUR especially seem to co-move over the mentioned time period. Of the four exchange rates mentioned above, the JPY is the only one that exhibits a slightly different movement.

[33]The given models are equivalent to the hypothesis formulated in equ. (20).
[34]The abbreviations refer to the USD expressed in units of foreign currency.

5.2 Unit Root Tests

Before the cointegration analysis can be carried out, it must be first controlled for the statistical requirements for the existence of a cointegration relationship. On this account, unit root tests are applied to all exchange rates in natural logarithm. In doing so, we expect that the order of integration is one because many studies dealing with the foreign exchange market could show a unit root in levels (e.g. Meese and Singelton (1982)). For this purpose, three unit root tests are taken. The Phillips-Perron (PP) test, the KPSS-test by Kwiatkowski et al. (1992), and the DF-GLS test by Elliott et al. (1996) are applied to the exchange rates in levels and in first differences.[35] The critical values for the PP test are taken from MacKinnon (1991).

The results of the unit root tests are depicted in table 1. Both the PP and the DF-GLS test are not able to reject the null hypotheses of a unit root in levels. In addition, the PP test rejects the null hypothesis of a unit root in first differences in all cases. With the exception of the EUR exchange rate for first differences the KPSS test confirms the results of the PP test. Contrary to the PP test and the KPSS test, the DF-GLS fails to reject the null hypothesis of non-stationarity for JPY, CAD and SEK but is able to reject a unit root for the EUR in first differences at the 5% significance level. Since the KPSS test is not able to reject stationarity for first differences with a small test statistic for the exchange rates JPY, CAD and SEK, all exchange rates are treated as first difference stationary, that is to say, the levels only contain a unit root. Based on this result, a cointegration analysis can be applied to the exchange rates.

5.3 Bivariate Cointegration Analysis

The results of the bivariate cointegration analysis are presented in table 2.[36] As can be seen, the null hypothesis of no cointegration can be rejected for model 1 for the exchange rate pairs EUR-GBP, EUR-AUD, GBP-CHF, GBP-AUD and GBP-SEK at the 5% level. Only in the case of the EUR-GBP pair does the rejection take place for the trace statistic. For model 2, only one rejection occurs for the pair CHF-AUD, again at the 5% level for both statistics. As far as rejection is concerned, model 3 basically shows the same results with the exception of EUR-CHF. Here, the null hypothesis can be rejected at the 5% level for both statistics. In the cases of the EUR-GBP and EUR-AUD, the rejection is significant at the 1% level. For all other exchange rate combinations, the null hypothesis of no cointegration cannot be rejected, at least at the 5% level. Model 1 (model 3) can show five (six) rejections of the null hypothesis of no cointegration, whereas model 2 only exhibits one. Contrary to earlier findings, a model with the inclusion of a constant (meaning either intercept or a linear trend in levels) as preferred by Diebold et al. (1994), Baillie and Bollerslev (1994), and Barkoulas and Baum (1997) can reject no cointegration.

For all pairings, heteroscedasticity cannot be rejected, i.e. significant ARCH effects are present. As Gonzalo (1994) and Rahbek et al. (2002) demonstrate, the Johansen procedure is quite robust in the case of heteroscedasticity while Lee and Tse (1996) show that ARCH effects have the tendency to overreject the null hypothesis of no cointegration but within a small magnitude. Although Che-

[35] The KPSS test and the DF-GLS test take the fact that the traditional (augmented) Dickey-Fuller test (ADF) test sometimes performs poorly into account. The DF-GLS test is more efficient than the traditional ADF test whereas the KPSS has the opposite formulation of the null hypothesis (stationarity vs. non-stationarity).

[36] The lag selection for the VAR is predominantly based upon the Schwartz-information criterion (SIC). See Reimers (1992), Cheung/Lai (1993), Ho/Sorensen (1996) and Gonzalo/Pitarakis (1998) for the reasons. Since the null hypothesis of no cointegration is more frequently rejected if the VAR is misspecified because of a lag length that is too short, it is only necessary to examine these results more closely where the null hypothesis of no cointegration could be rejected. In these cases, the lag length is adjusted by eliminating serial correlations. As long as serial correlation is still present in the residuals, the lag length is increased. To determine serial correlation, a LM test is taken.

ung and Lai (1993) suggest primarily using the trace statistic in these cases, we also rely on the maximum eigenvalue statistic. The reason is that Johansen and Juselius (1990) generally prefer to use it because of the precise distinction between the hypotheses. Taking the critique of Cheung and Lai (1993) and the findings of Lee and Tse (1996) into account, the significance level of 10% is omitted. Bearing in mind the problems with heteroscedasticity, the rejection at the 5% significance level should be firm evidence in favour of cointegration.

In order to evaluate the market efficiency, the corresponding VECM of each pair has to be estimated. However, before the results are given, the correct model has to be filtered out. Therefore, a likelihood-ratio test, as proposed by Johansen und Juselius (1990) and Johansen (1994), is carried out. The $LR-$test is a nested model test, meaning that it is investigated if model 1 is included in model 3. Thus, the $LR-$test is contemporaneously a test for the absence of a linear trend. If model 1 is included in model 3, the absence of a linear trend cannot be rejected, i.e. there is evidence against a linear trend in levels. The test statistic is

$$-2 \ln Q(H_1(r)|H_3(r)) = -T \sum_{i=r+1}^{p} \ln\{(1-\hat{\lambda}_i^1)/(1-\hat{\lambda}_i^3)\}. \tag{25}$$

and it is asymptotically distributed as χ^2 with $(p-r)$ degrees of freedom (Johansen, 1994, p. 213 and p. 217).

The results are presented in table 3. As can be seen, the null hypothesis cannot be rejected for all currency pairs. In this regard, model 1 represents the data more accurately.

Similarly, a test can be conducted to test the absence of the constant term in the CIV. This test is carried out in the same fashion as the previous one. Again, a $LR-$test can be used where it is being tested if model 2 is included in model 1. A rejection shows evidence in favour of a significant constant in the cointegration vector and in favour of non-zero mean. The test statistic is given

$$-2 \ln Q(H_2(r)|H_1(r)) = T \sum_{i=1}^{r} \ln\{(1-\hat{\lambda}_i^2)/(1-\lambda_i^1)\} \tag{26}$$

and it is asymptotically distributed as χ^2 with r degrees of freedom.

Examining the results depicted in table 3 reveals that a non-zero mean cannot be neglected and Model 1 and model 2 are not equivalent. Following Diebold et al. (1994), model 1 should be preferred, as should the correct critical values as emphasized by Johansen (1995) (p. 163). Returning to the cointegration test, it seems obvious that five cointegration relations exist in the bivariate investigation.

Based on the outlined findings, stronger co-movements seem to prevail amongst the pairs EUR, GBP, AUD, CHF and SEK. For the JPY, the null hypothesis is broadly rejected. The estimation results will be interpreted later. Table 4 presents the estimated μ coefficients and both the adjustment coefficients α and the cointegration parameters β. As presented in column 2, all constant terms within the error correction term are statistically different from zero. All cointegration parameters are statistically speaking significantly different from zero at the 1% level (column 4) and have the correct sign implied by the no arbitrage condition. In columns 5 and 8 for model 2 respectively, the no arbitrage condition is tested directly. This is done by restricting the cointegration vector corresponding to the no arbitrage condition. Therefore, the cointegration vector is linearly restricted so that

$$H_\beta : \beta = H\phi \tag{27}$$

where H is a $p \times s$ matrix of known parameter with s the number of unrestricted cointegration parameters. Hence, it is assumed that $r \leq s \leq p$, such as in the case of $r = s$ where the cointegration space is completely specified and in the case of $r = p$ where the cointegration space is free of restrictions. The unknown parameters in the $s \times r$ matrix ϕ has to be estimated. H is assumed to be orthogonal to a matrix R, so that $H = R_{\perp}$ and hence

$$H_{\beta} : R'\beta = 0. \tag{28}$$

The unrestricted cointegration vector in equ. (17) is replaced by ϕ such that

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \alpha\phi' X_{t-1} + \mu + \Phi D_t + \epsilon_t \tag{29}$$

results. Equ. (29) is used to proceed with the usual Johansen procedure (Johansen, 1988, 1991). Having solved the eigenvalue problem, the restricted model can be tested against the unrestricted one. To test market efficiency, the restrictions from equ. (15) enter the matrix H. The explicit test for market efficiency is thus a $LR-$test that becomes

$$-2\ln Q(H_{\beta}(r)|H(r)) = -T\sum_{i=1}^{r} \ln\{(1 - \hat{\lambda}_i^{\beta})/(1 - \lambda_i)\} \tag{30}$$

and is asymptotically distributed as χ^2 with $r(p - s)$ degrees of freedom (Johansen, 1995, p. 107). The non-rejection of the null hypothesis supports market efficiency. However, this can only be verified for the EUR-AUD and GBP-CHF pairings, the latter of which only with a p-value of 0.052. For all of the others, the correct cointegration vector is rejected. In addition, all constant terms are significantly different from zero with regard to the intercept. All constant terms are positive except that for the EUR-GBP pair. Here it is negative.

The adjustment coefficients are listed in column 3. Not all of them are statistically significant and they also partly have the wrong sign. Usually, in a correctly specified VECM, the α coefficient of the variable which stands in the first line in the cointegration vector must have a negative sign. All succeeding adjustment parameters should have a positive sign. The reason is that if the first variable exceeds its value implied by the long-run relationship, an adjustment to the long-run equilibrium occurs in following periods. Hence, it has to decrease in order to converge to the equilibrium relationship. Equally, the second variable has to increase and, thus, the sign of the adjustment coefficient has to be positive.

Only two currency pairs exhibit the correct signs (EUR-CHF and GBP-AUD). However, besides the sign, the significance of the adjustment coefficient is also important. If the adjustment coefficient is not statistically different from zero, the corresponding endogenous variable will not participate in the adjustment process. Thus, the complete adjustment to long-run equilibrium runs via the second endogenous variable. Engle et al. (1983) introduced the concept of weak exogeneity in this context. For a pair of currencies, the weak exogeneity of the first one would mean that the second currency carries the burden of the whole adjustment process. Johansen (1992b, 1992c) examines these cases closely. He establishes that weak exogeneity can be tested using simple $t-$tests if only one cointegration relation is present (needless to say that this is the maximum number of cointegration vectors in a bivariate setting). Alternatively, he proposes a likelihood-ratio test that uses a restriction of the corresponding adjustment coefficients very similar to the test for a restricted cointegration space and tests the restricted model against the original one. As is shown by Johansen and Juselius (1990), the α matrix is linearly restricted by matrix A with A as a $p \times s$ matrix of known parameters. The hypothesis is formulated as

$$H_\alpha : \alpha = A\psi \tag{31}$$

and with ψ a $s \times r$ matrix of unknown parameters that shall be estimated. s is the number of the remaining adjustment coefficients, i.e. the number of unrestricted coefficients or equivalently, the rank of the matrix A, so that m is the number of restrictions and r the rank of β with $s \geq r$. This means, the number of unrestricted coefficients must be greater than or equal to the number of cointegration vectors. Alternatively, the hypothesis H_α can also expressed as

$$H_\alpha : R'\alpha = 0. \tag{32}$$

In this formulation, the idea behind the test becomes much clearer. R is a $p \times (p - s)$ matrix that satisfies the condition $R = A_\perp$, i.e. R is orthogonal to A such that $R'A = 0$. It consists of the number of restrictions that restricts α to zero. The unrestricted $\alpha-$matrix in equ. (17) is replaced by ψ so that

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \psi\beta' X_{t-1} + \mu + \Phi D_t + \epsilon_t \tag{33}$$

results. Equ. (33) is used to proceed with the usual Johansen procedure (Johansen, 1995, pp. 124-127). With the results of this procedure at hand, the restricted model can be tested against the unrestricted one. The likelihood ratio test therefore becomes

$$-2 \ln Q(H_\alpha(r)|H(r)) = -T \sum_{i=1}^{r} \ln\{(1 - \hat{\lambda}_i^\alpha)/(1 - \lambda_i)\} \tag{34}$$

and is asymptotically distributed as χ^2 with $r(p - s)$ degrees of freedom (Johansen/Juselius, 1990, p. 200) or with the number of weakly exogenous variables with rm (Juselius, 2006, p. 195). A failure to reject the null hypothesis that the restricted model is included within the unrestricted model supports weak exogeneity.

In table 4, the asterisks refer to the usual $t-$test. Table 5 presents the results of the $LR-$test. The results of both types of tests coincide. After taking weak exogeneity into account, only the foreign exchange rate pairing EUR-GBP and GBP-CHF exhibits the wrong sign for the adjustment coefficient referring to the EUR and the CHF respectively. It should be mentioned that the hypothesis of weak exogeneity can only be rejected at the 5% level for the EUR contrary to the 1% level in all other cases.[37] The CHF for the pairing GBP-CHF is weakly exogenous.

Summarizing the results reveals that for the pairs EUR-CHF the CHF, EUR-AUD the EUR, the GBP-AUD the AUD, the GBP-SEK the SEK, and CHF-AUD the CHF are weakly exogenous. But what does that mean economically with respect to the long-run relationship and the adjustment between the exchange rates? This will be explained using the EUR-AUD pairing where the EUR is weakly exogenous. If the Euro appreciated (a fall in s_t), the Australian Dollar would appreciate as well in order to adjust to the long-run equilibrium. Thus, an appreciation of the Euro would cause an appreciation of the Australian Dollar. Alternatively, an appreciation of the Australian Dollar without a change in the EUR would cause the Australian Dollar to depreciate again. Although the cointegration vector fulfils the no arbitrage condition, the AUD reacts to changes in the EUR. Consequently, the EUR and the AUD are not different assets.

[37] Except the EUR-CHF where the CHF- $\alpha-$coefficient is not significant and the EUR $\alpha-$coefficient only at the 5% level

In addition to that, the results based on the Johansen approach imply that the CHF is weakly exogenous in the pairing with the EUR. The result should be challenged by examining the turnover of the EUR on the international foreign exchange market. In 2004, the EUR had a share of 28% compared to a share of 4% for the CHF (BIS, 2005, p. 10). The same argument can be extended to the pairings GBP-AUD and GBP-SEK. Generally, the GBP seems to be endogenous whereas the second exchange rate weakly exogenous. Since the GBP has a higher share in the global exchange market than the AUD, CHF and SEK, these results should be the subject of further investigations. The same can be said of the wrong sign of the adjustment coefficients for the pairings EUR-GBP and GBP-CHF.

5.4 Cointegration and Structural Breaks

Since the evidence in favour of cointegration (or in favour of the rejection of no cointegration) is far from striking, the robustness of the results must be verified. For this reason, an additional cointegration test that explicitly takes a structural change into account is provided. The test developed by Gregory and Hansen (1996) (hereafter GH) stands in the tradition of residual-based cointegration tests in the Engle/Granger mould. As described in an earlier section, the Engle/Granger approach uses residuals from the regression of one variable to the other one to test for unit roots. Stationary residuals are consistent with cointegration. The GH test accounts for one structural change that occurs at an unknown time. For the investigation, only two of the three models proposed by GH shall be used here with regard to the previous considerations concerning the market efficiency hypothesis. The estimation equations are presented below:

Model *level shift*:

$$y_t = \mu_1 + \mu_2 \phi_{t\tau} + \beta x_t + e_t \tag{35}$$

Model *regime shift*:

$$y_t = \mu_1 + \mu_2 \phi_{t\tau} + \beta_1 x_t + \beta_2 x_t \phi_{t\tau} + e_t \tag{36}$$

with

$$\phi_{t\tau} = \begin{cases} 0 & \text{if } t \leq [n\tau] \\ 1 & \text{if } t > [n\tau] \end{cases} \tag{37}$$

where τ is unknown and expresses the occurrence of the structural break (Gregory/Hansen, 1996, pp. 102/103).[38] GH start with a simple regression equation where a constant term (μ_1) is included similar to the one in the EG-approach and given in equ. (1). In the first model, the constant term is allowed to change at an unknown time ($\mu_2 \phi_{t\tau}$). According to GH this model is labelled "level shift". Economically, the change in the constant term would mean, that the mean of the cross rate have abruptly changed. Referring to the models outlined in the previous section, this model corresponds to model 1. In the second model provided by GH, besides the constant term the cointegration coefficient (β_1) is also allowed to change. The expression $\beta_2 x_t$ reflects the change in β_1. This model corresponds to the regime shift model in their contribution.[39] GH suggest estimating the model and checking for unit roots in the error term whilst taking into account a dummy variable. Therefore,

[38] The parameter $\tau \in (0, 1)$ can be seen as the timing of the relative change point, as Gregory/Hansen (1996) write.
[39] Besides these two models, GH also proposed a model with a linear trend where the constant term varies. This setting should be neglected due to the linear trend in the cointegration equation.

the dummy variable is allowed to vary successively over a specific range. For each potential break point, the cointegration test statistic is computed and the point in time is selected when the test statistic achieves its minimum value. This point in time determines the structural change and the test statistic is compared with the critical values as proposed by GH that are based upon the ADF critical values. Similarly, as in the EG approach, if the test statistic is below the critical value, the null hypothesis of no cointegration can be rejected in favour of the hypothesis cointegration with one endogenous regime shift.

Since the same problems arise, applying the GH test with respect to the choice of the correct dependent and independent variables as in the EG case, the test is carried out with an alternative setting. The results are displayed in table 6. As can be seen, the null hypothesis of no cointegration can again only be rejected for six pairings. Confirming the weak evidence in favour of the rejection of no cointegration by applying the Johansen procedure, this is not the case for the pairings GBP-CHF and GBP-AUD. There is only a weak rejection for the EUR-CHF pair in one setting. Thus, it can be concluded that cointegration is not present in the aforementioned pairs, even taking a break in the constant term or the cointegration parameter into account. A strong rejection of the null hypothesis can be considered for the pairing GBP and SEK both in the case of a break in the intercept and in the cointegration parameter. For both cases, a break is estimated to occur at the end of 2003 or the beginning of 2004. Since the results do not change in the second case, we can conclude that only a change in the constant term is important. For the pairing EUR-AUD, both specifications yield a rejection of the null hypothesis. However, the dates of the breakpoint differ significantly. Even the estimated breakpoints do not coincide in the level shift model. It can be concluded that a change in the cointegration parameter is important. With respect to the findings of the Johansen approach, the result for the pairing GBP and EUR is more interesting, especially as both currencies rank amongst the most important ones. While no cointegration against the alternative of cointegration in the face of a break in the intercept can only be rejected at the 5% level for the case where the EUR is the independent variable, evidence in favour of cointegration with breaks both in levels and the cointegration parameter is much more striking. The potential break dates back to the beginning of 2003. Contrary to the Johansen approach, the GH approach is able to reject the null hypothesis for the pairing EUR and SEK with a similar date for the structural break in both settings. Here, the breakpoint is estimated to occur at the end of 2000 and the beginning of 2001. For the pairings EUR-CHF and CHF-AUD, the null hypothesis of no cointegration can only be rejected in one case for each pair. Since the evidence in favour of the rejection based on the 5% level is weak, it is seen evidence against cointegration. Finally, the GH approach reveals that the null hypothesis of no cointegration can be rejected for four pairings. Since the estimation of the VECM for the pairing EUR-AUD has already shown the consistency with EMH, two sub-periods have to be investigated more carefully for the remaining pairings, namely the EUR-GBP, EUR-SEK and GBP-SEK.

5.5 Sub-periods for the EUR-GBP, EUR-SEK and the GBP-SEK pairings

Based on the estimated breakpoint of the GH approach, it is assumed that the break for the EUR-GBP pair occurred at the end of 2002 or at the beginning of 2003, for the GBP-SEK at the beginning of 2004 and for the EUR-SEK at the beginning of 2001. Hence, the complete samples are divided into two sub-samples. Beginning with the EUR-GBP, the first starts on 4 January 1999 and ends on 31 December 2002. Consistently, the second sub-sample runs from 3 January 2003 to 31 December 2006. The results of the complete cointegration analysis and VECM estimation for the EUR-GBP pairing are summarized in table 7. As we can see, the null hypothesis of no cointegration can be rejected at the 5% level in the first period for model 1 and model 3.[40] For the second period, the rejection of the null hypothesis of no cointegration is robust at the 1% level for model 1 and model 3. Surprisingly, the hypothesis of at least one cointegration vector is also rejected for model 3 at the 5% level.[41] For both periods, $LR-$tests are carried out to find the correct model. Similar to the results of the complete period, model 1 is the correct one in comparison to model 3. Therefore, the results for model 3 can be neglected. Again, the constants are both significantly different from zero and negative and the cointegration parameter has the correct signs. As we can see very clearly, the absolute value of the cointegration parameter for the GBP is greater than one in the first period and less than 1 in the second period, confirming the evidence for a structural break indicated by the GH test. Nevertheless, the restricted model with respect to the cointegration vector cannot be rejected marginally for both periods; for the first period with a p-value of 0.054 and for the second period with a p-value of 0.067. The magnitude of the constant terms and hence the mean of the cross rate is halved from the first to the second period reflecting a level shift of the cross rate. Contrary to the estimation of the whole period, in both sub-samples the adjustment coefficient for the EUR is not statistically different from zero and the corresponding null for the GBP can be rejected very clearly at the 1% level. The $LR-$test for weak exogeneity supports these results. It can be shown that for both sub-periods the EUR is weakly exogenous, although the restricted α_{GBP} can only be rejected at the 5% level. The speed of adjustment is higher in the second period than in the first. 50% of a shock to long-run equilibrium is cancelled out after 7.5 weeks in the first period and 5 weeks in the second period.[42] Appreciations of the Euro seem to cause the British Pound to appreciate as well. Alternatively, appreciations of the British Pound would in turn result in depreciations of the British Pound. In addition, the cointegration vector corresponds to the no arbitrage condition in both sub-samples but with weak evidence. Consequently, it can be stated that the market is efficient despite cointegration, although the evidence is weak.

The results of the closer investigation of the two sub-samples for the EUR-SEK pairing are given in table 8. Again, model 1, taking a non-zero mean into account, represents the model more accurately than model 3. Hence, the null hypothesis of no cointegration can only be rejected for the second sub-sample. For this period, the cointegration vector is consistent with EMH and the market can be regarded as efficient. Nevertheless, the results are very fragile because not the EUR but the SEK is weakly exogenous and the corresponding hypotheses regarding the EUR are only slightly rejected at the 5% level. In addition, the results would show that the most traded exchange rate in the world, namely the Euro-U.S. Dollar exchange rate, is influenced by the less important SEK. In addition, the adjustment speed is lower than for the paring EUR-GBP. The results regarding the EUR-SEK pairing should be challenged.

A similar argumentation can be applied to the GBP-SEK pairing. A closer look at the two sub-periods for the GBP-SEK pairing depicted in table 9 yields a broad non-rejection of the null hypothesis of no cointegration in the first period. Complementary to the EUR-SEK pairing, the null hypothesis can only be rejected for the second sub-sample and only for models 1 and 3. As we can see from examining the relevant $LR-$test, model 1 represents the data the most accurately. The cointegration vector has the correct signs but the hypothesis of the restriction of the cointegration space with respect to EMH can be rejected. As a result, it can be stated that for the GBP-SEK

[40] Here, the rejection takes place at the 1% level for the trace statistic.

[41] For model 2 the null hypothesis can be rejected in the second period but only at the 5% level. Since the levels of the EUR and the GBP do not coincide, model 2 can be neglected.

[42] Equation to apply: $t = \frac{\ln(1-x)}{\ln(1-\alpha)}$, with α as the adjustment coefficient and t as the time required for a shock to dissipate by x percent.

pairing EMH does not hold in the second period. Here, the SEK seems to be weakly exogenous. The $t-$statistic for the GBP is significant at the 1% level and the corresponding $LR-$test of weak exogeneity provides a strong rejection of the hypothesis as can be seen from the p-value of nearly zero. The $LR-$test supports slightly weak exogeneity with respect to the SEK (with a p-value of 0.06) whereas the $t-$test rejects the hypothesis that the $\alpha_{SEK}-$ coefficient is significantly different from zero at the 5% level. Therefore, the negative sign is important and not consistent with the expected adjustment process.

Subsequently, the market efficiency should be questioned due to the weak exogeneity of the SEK in the cases of EUR-SEK and GBP-SEK because of the minor economic importance of the SEK. Regarding the weak exogeneity of the EUR for the pairing EUR-GBP, the results indicate that the EUR only acts independently for the last case. Bearing in mind the higher weight of the EUR and the GBP on the foreign exchange market, the results for the EUR-SEK and the GBP-SEK pairing are surprising. Therefore, finding cointegration based on the Johansen approach should be challenged and treated with caution for these two cases whereas the findings for the EUR-GBP pair are rather robust.

6 Conclusion

In this paper, the foreign exchange market is tested for cointegration between pairs of daily foreign exchange rates. The period of observation covers the introduction of the Euro and the most recent period of floating exchange rates thereafter. Motivated by an argument provided by Granger (1986), a market is not efficient if cointegration between pairs of exchange rates can be considered because the observation of cointegration means the predictability of at least one exchange rate.
The empirical analysis draws on the Johansen (1988, 1991) approach. It is shown that the null hypothesis can be rejected for most of the exchange rates. Only for seven exchange rate pairs can the hypothesis of no cointegration be rejected on the basis of weak evidence. A likelihood-ratio test for the correct model shows that, except for one pairing, a model specification with a constant term in the error correction term represents the best fit. These results partly contradict and clarify the results by Diebold et al. (1994), Barkoulas and Baum (1997), and many more. In the bivariate analysis, these results prove that a linear trend is not present. The estimation of the vector error correction model illustrates that the cointegration vector is only consistent with market efficiency in two cases. In all other cases, the rejection of the correct cointegration vector would imply an inefficient market. After the application of the Gregory-Hansen cointegration test, no cointegration can only be rejected for the pairings EUR-GBP, EUR-SEK, GBP-SEK and EUR-AUD. Since the estimation results of the Johansen approach for the EUR-GBP pair are not convincing with respect to the wrong sign of the adjustment parameters, the break point estimated by the Gregory-Hansen approach is used to generate two sub-samples. The results of the applied Johansen test and the estimated coefficients of the VECM for these two sub-samples are more striking than for the overall period of observation. It can be shown for the EUR-GBP paring that the rejection of no cointegration is much more robust for both periods (4 January 1999 to 31 December 2002 and 3 January 2003 to 31 December 2006). For both sub-periods market efficiency cannot be rejected. It is noteworthy that in both cases the EUR is weakly exogenous. Hence, the causality runs from EUR to GBP. Consequently, the EUR/USD and the GBP/USD exchange rates are not different assets whereas the EUR/USD exchange rate influences the GBP/USD exchange rate. Based on bivariate cointegration analysis, the conclusion can be drawn that the foreign exchange market is broadly consistent with market efficiency in the sense of Granger (1986). Only for a small number of exchange rate pairings can the null hypothesis of no cointegration be rejected. As a result, it can be considered that the introduction of the Euro has not resulted in an inefficient market. Information regarding the development of one exchange rate cannot be taken from the past realisations of a different exchange rate to earn excess returns, although the Australian Dollar, the British Pound Sterling, and the Swedish Krona seem to be linked to the Euro. The factors that have caused the cointegration relationship should be the subject of future research.

A Appendix

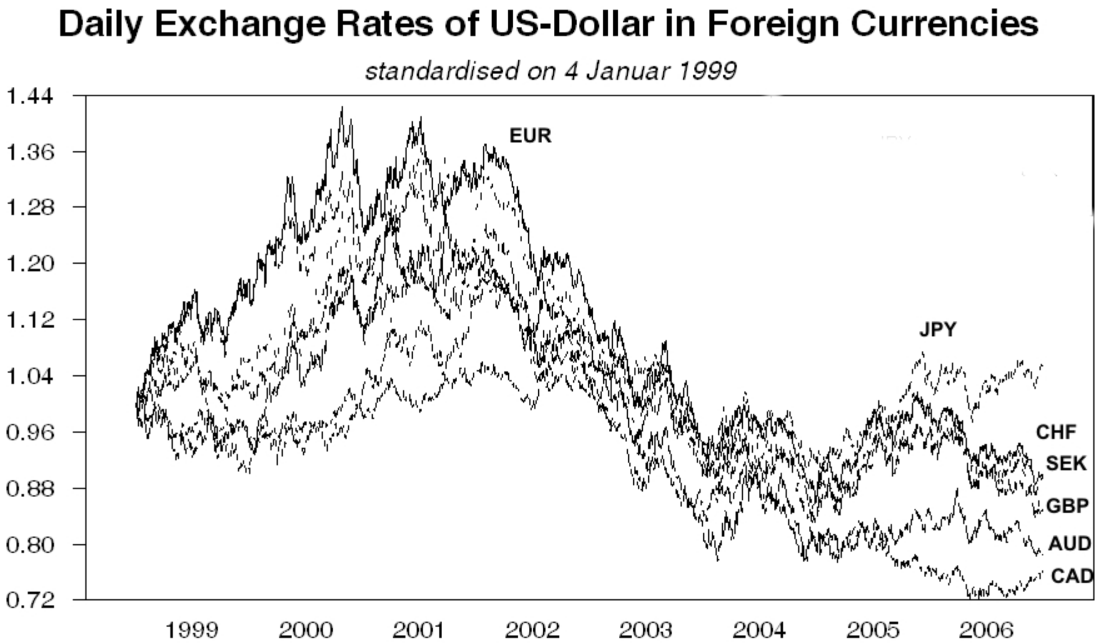

Figure 1: Daily exchange rates of the USD (USD in foreign currencies) standardised on the EUR/USD exchange rate on 4 January 1999

Source: Federal Reserve Bank of St. Louis (2007).

Table 1: Unit Root Tests for the USD exchange rates

Unit Root Tests for the USD exchange rates					
	PP test	**DF-GLS test**		**KPSS test**	
First differences	**Statistic**[a]	**Lags**	**Statistic**[b]	**Lags**	**Statistic**[c]
EUR/USD	-44.751**	25	-2.000*	11	0.466*
GBP/USD	-43.994**	25	-6.043**	41	0.276
JPY/USD	-46.103**	24	-1.123	19	0.062
CHF/USD	-46.352**	25	-4.117**	6	0.261
AUD/USD	-43.61**	25	-1.989*	29	0.221
CAD/USD	-44.763**	25	-1.83	34	0.162
SEK/USD	-44.300**	25	-1.716	39	0.309
Levels					
EUR/USD	-0.573	4	-0.731	30	4.52**
GBP/USD	-0.514	4	-0.595	30	4.46**
JPY/USD	-2.162	4	-1.928	29	0.65*
CHF/USD	-0.800	4	-0.839	30	5.07**
AUD/USD	-0.553	7	-0.411	30	4.06**
CAD/USD	-0.367	1	0.513	30	5.24**
SEK/USD	-0.588	1	-0.743	30	3.57**

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

Table 2: Test for bivariate cointegration using the Johansen approach

Johansen Cointegration Test[a]									
T=2011				**Model 1: constant in CI**		**Model 2: no constant**		**Model 3: linear trend in levels**	
USD		l[b]	$H_0 : r \leq$	λ_{trace}	λ_{max}	λ_{trace}	λ_{max}	λ_{trace}	λ_{max}
EUR									
	GBP	3	0	**21.1981***	**20.1295***	2.3595	2.1488	**20.6314****	**19.5908****
			1	1.0686	1.0686	0.2106	0.2106	1.0406	1.0406
	JPY	2	0	7.9216	7.3258	0.6793	0.6498	7.5837	7.3251
			1	0.5959	0.5959	0.0295	0.0295	0.2587	0.2587
	CHF	3	0	17.2939	14.6374	4.0165	3.9528	**17.0814***	**14.5754***
			1	2.6565	2.6565	0.0637	0.0637	2.506	2.506
	AUD	3	0	19.8491	**19.3666***	2.0315	1.5521	**19.2571***	**18.9967***
			1	0.4825	0.4825	0.4794	0.4794	0.2604	0.2831
	CAD	2	0	11.0619	10.1813	3.323	3.1001	9.3171	9.3141
			1	0.8806	0.8806	0.2229	0.2229	0.003	0.003
	SEK	2	0	6.9943	6.1989	1.1052	0.8614	6.7434	6.0854
			1	0.7954	0.7954	0.2438	0.2438	0.658	0.658
GBP									
	JPY	2	0	6.294	5.4315	1.222	0.8775	5.5427	5.406
			1	0.8624	0.8624	0.3447	0.3447	0.1367	0.1367
	CHF	2	0	18.2949	**15.7678***	3.2622	2.985	**17.7693***	**15.2471***
			1	2.5271	2.5271	0.2772	0.2772	2.5222	2.5222
	AUD	2	0	17.1242	**16.253***	1.0094	0.8945	**16.3642***	**16.2326***
			1	0.8712	0.8712	0.1149	0.1149	0.1316	0.1316
	CAD	2	0	9.9766	8.0597	2.1375	1.993	8.0827	8.0553
			1	1.9169	1.9169	0.1445	0.1445	0.0274	0.0274
	SEK	2	0	16.6973	**16.0294***	0.8908	0.725	**16.1741***	**15.7683***
			1	0.6679	0.6679	0.1658	0.1658	0.4058	0.4058
JPY									
	CHF	2	0	6.958	6.1052	0.9355	0.9037	6.6186	6.1047
			1	0.8527	0.8527	0.0318	0.0318	0.5139	0.5139
	AUD	2	0	9.5237	8.2631	1.2908	1.2892	8.6917	8.2457
			1	1.2606	1.2606	0.0016	0.0016	0.4459	0.4459
	CAD	2	0	7.4805	5.3118	2.1834	2.1712	5.4641	5.3028
			1	2.1686	2.1686	0.0123	0.0123	0.1613	0.1613
	SEK	2	0	11.8501	10.8277	1.319	1.1423	11.4192	10.8202
			1	1.0224	1.0224	0.1766	0.1766	0.599	0.599
CHF									
	AUD	4	0	14.1742	13.1891	**13.6315***	**13.1803***	13.5534	12.7338
			1	0.9852	0.9852	0.4512	0.4512	0.8196	0.8196
	CAD	2	0	9.6006	8.8564	9.0747	8.3486	7.8541	7.7306
			1	0.7442	0.7442	0.7261	0.7261	0.1235	0.1235
	SEK	2	0	8.914	6.9663	3.3578	3.0515	8.6466	6.7654
			1	1.9477	1.9477	0.3063	0.3063	1.8812	1.8812
AUD									
	CAD	2	0	7.5157	6.1153	7.0242	5.6241	5.7541	5.6611
			1	1.4003	1.4003	1.4001	1.4001	0.093	0.093
	SEK	2	0	19.7392	19.134	1.1876	1.1275	19.1403	18.9276
			1	0.6052	0.6052	0.0601	0.0601	0.2127	0.2127
CAD									
	SEK	2	0	8.3748	7.2695	2.3413	1.983	6.6244	6.6069
			1	1.1053	1.1053	0.3583	0.3583	0.0175	0.0175

$*$ Rejection of the nullhypothesis at the 5% significants level.
$**$ Rejection of the nullhypothesis at the 1% significants level.
a Critical values from Johansen/Juselius (1990) and Osterwald-Lenum (1992).
b Determination of the lag length (l) based on SIC and in the face of remaining serial correlation on the basis of the Lagrange-Multiplier (LM) tests for autocorrelation in the residuals of the underlying VAR.
λ_{trace} refers to the trace statistic and λ_{max} refers to the maximum eigenvalue statistic. see Johansen (1988, 1991).

Table 3: Test on model specification of the VECM

		Test on model Specification[a]	
		Test on the absence of a linear trend	**Test on the absence of the constant in the CIV**
		Model 1 in Model 3	**Model 2 in Model 1**
		$-2\ln Q(H_1(r)\|H_3(r)) = -T\sum_{i=r+1}^{p}\ln\left(\frac{(1-\hat{\lambda}_i^1)}{(1-\hat{\lambda}_i^3)}\right)$	$-2\ln Q(H_2(r)\|H_1(r)) = T\sum_{i=1}^{r}\ln\left(\frac{(1-\hat{\lambda}_i^2)}{(1-\hat{\lambda}_i^1)}\right)$
EUR	GBP	0.02	**17.981****
EUR	CHF	0.14	**10.685****
EUR	AUD	0.22	**17.815****
GBP	CHF	0.02	**12.783****
GBP	AUD	0.723	**15.358****
GBP	SEK	0.26	**15.304****

$*$ Rejection of the null hypothesis at the 5% significance level.
$**$ Rejection of the null hypothesis at the 1% significance level.
a Critical values 3.84 for 5% and 6.63 for 1% significance level.

Table 4: Estimation of the VECM

Estimation of the VECM: Adjustmentcoefficients and Cointegration Vector							
	Model 1: non-zero mean				**Model 2: no constant**		
	μ_0	α	β	$LR(1,-1)^a$	α	β	$LR(1,-1)^a$
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
EUR		**0.009***	1	**11.369****			
GBP	-0.628**	**0.015****	-1.405**	(0.001)			
EUR		**-0.009***	*1*	**5.512***			
CHF	*0.494***	*0.005*	*-1.17***	(0.019)			
EUR		*0.001*	*1*	1.007			
AUD	*0.481***	**0.014****	*-0.93***	(0.316)			
GBP		**-0.013****	*1*	3.765			
CHF	*0.781***	**-0.009***	*-0.781***	(0.052)			
GBP		**-0.011****	*1*	**9.616****			
AUD	*0.796***	*0.007*	*-0.673***	(0.002)			
GBP	*2.111***	**-0.013****	*1*	**6.849****			
SEK		*-0.001*	*-0.757***	(0.009)			
CHF					0.0012	1	**6.707****
AUD					**0.009****	-0.829**	(0.01)

$*$ Rejection of the null hypothesis at the 5% significance level.
$**$ Rejection of the null hypothesis at the 1% significance level.
[a] Test statistic for the hypothesis of a restricted cointegration vector. P-values in brackets.
μ_0 represents the intercept in the error correction term. the coefficients α and β
refer to the adjustment coefficient and the cointegration parameter respectively.

Table 5: Test for weak exogeneity

Test for weak exogeneity[a]			
	Model 1	**Model 2**	**Model 3**
	$LR(\alpha_i)$	$LR(\alpha_i)$	$LR(\alpha_i)$
EUR	**4.38***		**4.161***
GBP	**17.558****		**17.022****
EUR	**3.901***		3.839
CHF	*0.835*		0.793
EUR	*0.011*		0.001
AUD	**14.669****		**14.314****
GBP	**12.51****		**12.00****
CHF	*3.568*		3.384
GBP	**7.737****		**7.973****
AUD	*1.574*		1.725
GBP	**9.96****		**9.709****
SEK	*0.028*		0.01
CHF		0.255	
AUD		**11.385****	

$*$ Rejection of the null hypothesis at the 5% significance level.
$**$ Rejection of the null hypothesis at the 1% significance level.
a Critical values 3.84 for 5% and 6.63 for 1% significance level.
The statistic $LR(\alpha_i)$ refers to the adjustment coefficient
in each line that is restricted to zero.

Table 6: Test for structural breaks - Gregory/Hansen (1996) cointegration test

			Break in Intercept		Break in Intercept + Cointegration parameter	
		Lags[a]	test statistic[b]	Breakpoint	test statistic[c]	Breakpoint
EUR	GBP	4	**-4.778***	27 January 2003	**-5.492****	11 March 2003
GBP	EUR	4	-4.293	20 January 2004	**-5.038***	17 March 2003
EUR	JPY	1	-4.023	19 October 2005	-4.023	19 October 2005
JPY	EUR	0	-4.023	19 October 2005	-4.023	19 October 2005
EUR	CHF	0	-4.513	24 September 2001	-4.616	24 September 2001
CHF	EUR	0	**-4.755***	24 September 2001	-4.805	24 September 2001
EUR	AUD	3	**-5.056***	24 April 2002	**-5.051***	24 April 2002
AUD	EUR	3	**-4.718***	04 January 2001	-4.771	23 October 2000
EUR	CAD	0	-4.023	19 October 2005	-4.243	29 July 2002
CAD	EUR	0	-4.096	17 June 2005	-4.105	30 June 2005
EUR	SEK	3	**-4.797***	09 January 2001	**-5.427***	06 December 2000
SEK	EUR	3	**-4.712***	09 January 2001	**-5.442***	06 December 2000
GBP	JPY	3	-4.023	19 October 2005	-4.023	19 October 2005
JPY	GBP	0	-4.023	19 October 2005	-4.023	19 October 2005
GBP	CHF	4	-3.695	28 September 2001	-3.566	29 August 2001
CHF	GBP	4	-4.361	11 June 2002	-4.363	11 June 2002
GBP	AUD	0	-4.337	16 January 2004	-4.474	26 November 2004
AUD	GBP	0	-4.337	02 October 2000	-4.800	02 September 2003
GBP	CAD	0	3.491	24 June 2005	-3.763	06 August 2002
CAD	GBP	0	-3.820	01 July 2005	-4.100	15 July 2005
GBP	SEK	6	**-5.775****	22 December 2003	**-5.913****	22 December 2003
SEK	GBP	6	**-5.415****	22 December 2003	**-5.422***	16 January 2004
JPY	CHF	0	-3.974	19 October 2005	-3.974	19 October 2005
CHF	JPY	0	-3.974	19 October 2005	-3.974	19 October 2005
JPY	AUD	0	-3.974	19 October 2005	-3.974	19 October 2005
AUD	JPY	0	-3.974	19 October 2005	-3.974	19 October 2005
JPY	CAD	0	-3.974	19 October 2005	-3.834	02 January 2003
CAD	JPY	4	-3.174	21 July 2005	-4.178	11 September 2003
JPY	SEK	1	-2.870	19 October 2005	-3.588	27 November 2000
SEK	JPY	0	-2.870	19 October 2005	-2.871	19 October 2005
CHF	AUD	3	**-4.681***	29 May 2002	-4.712	29 May 2002
AUD	CHF	3	-4.017	05 April 2003	-4.010	20 November 2002
CHF	CAD	0	-4.186	25 June 2002	-4.239	05 July 2002
CAD	CHF	0	-3.541	01 July 2005	-3.657	17 September 2004
CHF	SEK	4	-4.512	12 June 2001	-4.544	01 August 2001
SEK	CHF	4	-4.238	12 June 2001	-4.330	12 June 2001
AUD	CAD	3	-3.254	19 July 2005	-3.532	22 November 2002
CAD	AUD	2	-3.632	19 July 2005	-4.485	08 October 2004
AUD	SEK	3	-4.253	18 April 2005	-4.810	29 June 2005
SEK	AUD	3	-4.215	18 April 2005	-4.534	28 October 2004
CAD	SEK	3	-3.789	16 May 2005	-3.958	25 July 2005
SEK	CAD	3	-3.378	29 June 2005	-3.986	01 August 2002

a The AIC is used for the lag selection.
b Critical values -5.13 for 1% significance level and -4.61 for 5% significance level.
c Critical values -5.47 for 1% significance level and -4.95 for 5% significance level.
$*$ Rejection of the null hypothesis at a significance level of 5%.
$**$ Rejection of the null hypothesis at a significance level of 1%.

Table 7: Estimation of subperiods for the paring EUR-GBP

		Estimation of subperiods for the pairing EUR-GBP					
		04 January 1999 - 31 December 2002			**03 January 2003 - 31 December 2006**		
		Model 1	**Model 2**	**Model 3**	**Model 1**	**Model 2**	**Model 3**
λ_{trace}	$r=0$	**21.545***	6.884	**21.289****	**30.352****	**13.53***	**28.653****
	$r<0$	5.670	0.019	**5.497***	7.792	0.913	**6.112***
λ_{max}	$r=0$	**15.867***	6.865	**15.793***	**22.560****	**12.617***	**22.54****
	$r=1$	5.678	0.019	5.497	73792.000	0.913	**6.112***
	Lags	5	5	5	2	2	2
α_{EUR}		0.001		0.002	-0.002	**-0.011****	-0.002
α_{GBP}		**0.017****		**0.017****	**0.027****	0.000	**0.027****
β_{EUR}		1		1	1	1	1
β_{GBP}		**-1.412****		**-1.416****	**-0.856****	**-0.356****	**0.855****
μ		**-0.637****			**-0.297****		
$LR(\alpha_{EUR})$		0.023		0.036	0.026	**5.106***	0.037
		(0.878)		(0.849)	(0.871)	(0.024)	(0.848)
$LR(\alpha_{GBP})$		**6.472***		**6.652***	**5.54***	0.000	**6.098***
		(0.011)		(0.01)	(0.019)	(0.993)	(0.014)
$LR((1,-1))$		3.702		3.806	3.356	**12.072****	3.518
		(0.054)		(0.051)	(0.067)	(0.001)	(0.061)
$LR(H_1(r)\|H_3(r))$		0.181			1.682		
$LR(H_2(r)\|H_1(r))$			**9.001****			**9.967****	
			(0.003)			(0.002)	

$*$ Rejection of the null hypothesis at the 5% significance level.
$**$ Rejection of the null hypothesis at the 1% significance level.
p-values in brackets.
λ_{trace} refers to the trace statistic and λ_{max} refers to the maximum eigenvalue statistic. see Johansen (1988, 1991)
$LR(\alpha_i)$ is a LR test on weak exgoneity of the adjustment coefficients.
$LR(1,-1)$ tests for a restricted cointegration vector.
$LR(H_i(r)|H_j(r))$ is the test statistic for testing model i against model j.

Table 8: Estimation of subperiods for the paring EUR-SEK

		Estimation of subperiods for the pairing EUR-SEK					
		04 January 1999 - 29 December 2000			**02 January 2001 - 31 December 2006**		
		Modell 1	**Modell 2**	**Modell 3**	**Modell 1**	**Modell 2**	**Modell 3**
λ_{trace}	$r=0$	19.269	11.165	**16.810***	**23.234***	2.705	**21.085****
	$r<0$	7.539	0.963	**5.331***	2.031	0.019	0.379
λ_{max}	$r=0$	11.730	10.202	11.479	**21.203****	2.686	**20.706****
	$r=1$	7.539	0.963	**5.331***	2.031	0.019	0.379
	Lags	1	1	1	1	1	1
α_{EUR}					**-0.020***		**-0.018***
α_{SEK}					0.002		0.004
β_{EUR}					1		1
β_{SEK}					**-0.978****		**-0.978****
μ					**2.174****		
$LR(\alpha_{EUR})$					**3.999***		3.583
					(0.046)		(0.058)
$LR(\alpha_{SEK})$					0.032		0.144
					(0.857)		(0.704)
$LR((1,-1))$					0.741		0.698
					(0.389)		(0.404)
$LR(H_1(r)\|H_3(r))$		2.211			1.659		
$LR(H_2(r)\|H_1(r))$			1.528			**17.661****	
			(0.216)			**(0.000)**	

$*$ Rejection of the null hypothesis at the 5% significance level.
$**$ Rejection of the null hypothesis at the 1% significance level.
p-values in brackets.
λ_{trace} refers to the trace statistic and λ_{max} refers to the maximum eigenvalue statistic. see Johansen (1988, 1991)
$LR(\alpha_i)$ is a LR test on weak exgoneity of the adjustment coefficients.
$LR(1,-1)$ tests for a restricted cointegration vector.
$LR(H_i(r)|H_j(r))$ is the test statistic for testing model i against model j.

Table 9: Estimation of subperiods for the paring GBP-SEK

		Estimation of subperiods for the pairing GBP-SEK					
		04.01.1999 - 31 December 2003			02 January 2004 - 31 December 2006		
		Modell 1	**Modell 2**	**Modell 3**	**Modell 1**	**Modell 2**	**Modell 3**
λ_{trace}	$r = 0$	16.461	0.425	16.209	**26.143****	4.492	**25.735****
	$r < 0$	0.329	0.008	0.179	3.557	0.267	3.553
λ_{max}	$r = 0$	16.132	0.417	16.033	**22.586****	4.225	**22.182****
	$r = 1$	0.329	0.008	0.179	3.557	0.267	**3.553**
	Lags	7	7	7	5	5	5
α_{GBP}					**-0.058****		**-0.058****
α_{SEK}					**-0.037***		**-0.036***
β_{GBP}					1		1
β_{SEK}					**-0.657****		**-0.657****
μ					**1.921****		
$LR(\alpha_{GBP})$					**12.936****		**12.753****
					(0.000)		(0.000)
$LR(\alpha_{SEK})$					3.546		3.468
					(0.060)		(0.063)
$LR((1,-1))$					**7.377****		**7.319****
					(0.007)		(0.007)
$LR(H_1(r)\|H_3(r))$		0.149			0.076		
$LR(H_2(r)\|H_1(r))$			**15.715****			**14.129****	
			(0.000)			(0.000)	

$*$ Rejection of the null hypothesis at the 5% significance level.
$**$ Rejection of the null hypothesis at the 1% significance level.
p-values in brackets.
λ_{trace} refers to the trace statistic and λ_{max} refers to the maximum eigenvalue statistic. see Johansen (1988, 1991)
$LR(\alpha_i)$ is a LR test on weak exgoneity of the adjustment coefficients.
$LR(1,-1)$ tests for a restricted cointegration vector.
$LR(H_i(r)|H_j(r))$ is the test statistic for testing model i against model j.

References

[1] Aroskar, R., S. K. Sarkar and P. E. Swanson (2004). European Foreign Exchange Market Efficiency: Evidence based on Crisis and Noncrisis Periods. *International Reviews of Financial Analysis*, Vol. 13:333–347.

[2] Baffes, J. (1994). Does Comovement among Exchange Rates imply Market Efficiency? *Economics Letters*, Vol. 44:273–280.

[3] Baillie, R. T. and T. Bollerslev (1989). Common Stochastic Trends in a System of Exchange Rates. *The Journal of Finance*, Vol. 44:167–181.

[4] Baillie, R. T. and T. Bollerslev (1994). Cointegration, Fractional Cointegration, and Exchange Rate Dynamics. *The Journal of Finance*, Vol. 49:737–745.

[5] Bank for International Settlements (2005). *Triennial Central Bank Survey: Foreign Exchange and Derivatives Market Activity in 2004*. Basel.

[6] Barkoulas, J. and C. F. Baum (1997). A Re-examination of the fragility of Evidence from Cointegration-Based Tests of Foreign Exchange Market Efficiency. *Applied Financial Economics*, Vol. 7:635–643.

[7] Barkoulas, J., C. F. Baum and A. Chakraborty (2003). Forward Premiums and Market Efficiency: Panel Unit-root Evidence from the Term Structure of Forward Premiums. *Journal of Macroeconomics*, Vol. 25:109–122.

[8] Cheung, Y.-W. and K. S. Lai (1993). Finite-Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration. *Oxford Bulletin of Economics and Statistics*, Vol. 55:313–329.

[9] Coleman, M. (1990). Cointegration-Based Tests of Daily Foreign Exchange Market Efficiency. *Economics Letters*, Vol. 32:53–59.

[10] Copeland, L. S. (1991). Cointegration Tests with Daily Exchange Rate Data. *Oxford Bulletin of Economics and Statistics*, Vol. 53:185–198.

[11] Crowder, W. J. (1994). Foreign Exchange Market Efficiency and Common Stochastic Trends. *Journal of International Money and Finance*, Vol. 13:551–564.

[12] Crowder, W. J. (1996). A Note on Cointegration and International Capital Market Efficiency: a Reply. *Journal of International Money and Finance*, Vol. 15:661–664.

[13] Diebold, F. X., J. Gardeazabal and K. Yilmaz (1994). On Cointegration and Exchange Rate Dynamics. *The Journal of Finance*, Vol. 49:727–735.

[14] Dwyer, Jr., G. P. and M. S. Wallace (1992). Cointegration and Market Efficency. *Journal of International Money and Finance*, Vol. 11:318–327.

[15] Elliott, G.; T. J. Rothenberg and J. H. Stock (1996). Efficient tests for an autoregressive unit root. *Econometrica*, Vol. 64:813–836.

[16] Engel, C. (1996). A Note on Cointegration and International Capital Market Efficiency. *Journal of International Money and Finance*, Vol. 15:657–660.

[17] Engle, R. F. and C.W.J. Granger (1987). Co-Integration and Error Correction: Representation, Estimation and Testing. *Econometrica*, Vol. 55:251–276.

[18] Engle, R. F. ; D. F. Hendry and J.-F. Richard (1983). Exogeneity. *Econometrica*, Vol. 51:277–304.

[19] Fama, E. F. (1965a). The Behaviour of Stock-Market Prices. *The Journal of Business*, Vol. 38:34–105.

[20] Fama, E. F. (1965b). Random Walks in Stock Market Prices. *Financial Analysts Journal*, Vol. 21:55–59.

[21] Fama, E. F. (1970). Efficient Capital Markets: A Review of Theory and Empirical Work. *The Journal of Finance*, Vol. 25:383–417.

[22] Fama, E. F. (1976). *Foundations of Finance: Portfolio Decisions and Securities Prices*. New York.

[23] Ferré, M. and S. G. Hall (2002). Foreign Exchange Market Efficiency and Cointegration. *Applied Financial Economics*, Vol. 12:131–139.

[24] Frenkel, J. A. and R. M. Levich (1975). Covered Interest Arbitrage: Unexploited Profits? *The Journal of Political Economy*, Vol. 83:325–338.

[25] Gonzalo, J. (1994). Five Alternative Methods of Estimating Long-run Equilibrium Relationships. *Journal of Econometrics*, Vol. 60:206–233.

[26] Gonzalo, J. and J.-Y. Pitarakis (1998). Specification via Model Selection in Vector Error Correction Models. *Economics Letters*, Vol. 60:321–328.

[27] Granger, C.W.J. (1981). Some Properties of Time Series Data and their Use in Econometric Model Specification. *Journal of Econometrics*, Vol. 16:121–130.

[28] Granger, C.W.J. (1983). Co-Integrated Variables and Error-Correcting Models. *UCSD Discussion Paper*, No. 83-13.

[29] Granger, C.W.J. (1986). Developments in the Study of Cointegrated Economic Variables. *Oxford Bulletin of Economics and Statistics*, Vol. 48:213–228.

[30] Granger, C. W.J. and N. Hyung (2004). Occasional Structural Breaks and Long Memory with an Application to the S&P 500 absolute stock returns. *Journal of Empirical Finance*, Vol. 11:399–421.

[31] Granger, C. W.J. and R. Joyeux (1980). An Introduction to long Memory Time Series Models and Fractional Differencing. *Journal of Times Series Analysis*, Vol. 1:15–39.

[32] Gregory, A. W. and B. E. Hansen (1996). Residual-based Tests for Cointegration in Models with Regime Shifts. *Journal of Econometrics*, Vol. 70:99–126.

[33] Hakkio, C. S. and M. Rush (1989). Market Efficiency and Cointegration: An Application to the Sterling and Deutschmark Exchange Markets. *Journal of International Money and Finance*, Vol. 8:75–88.

[34] Haug, A. A., J. G. MacKinnon and L. Michelis (2000). European Monetary Union: A Cointegration Analysis. *Journal of International Money and Finance*, Vol. 19:419–432.

[35] Ho, M. S. and B. E. Sorensen (1996). Finding Cointegration Rank in High Dimensional Systems Using the Johansen Test: An Illustration Using Data Based Monte Carlo Simulations. *The Review of Economics and Statistics*, Vol. 78:726–732.

[36] Jensen, M. C. (1978). Some Anomalous Evidence Regarding Market Efficiency. *Journal of Financial Economics*, Vol. 6:95–101.

[37] Jeon, B. N. and B. Seo (2003). The Impact of Asian Financial Crisis on Foreign Exchange Market Efficiency: The Case of East Asian Countries. *Pacific-Basin Finance Journal*, Vol. 11:509–525.

[38] Jeon, B. N. and E. Lee (2002). Foreign Exchange Market Efficiency, Cointegration, and Policy Coordination. *Applied Economics Letters*, Vol. 9:61–68.

[39] Johansen, S. (1988). Statistical Analysis of Cointegration Vectors. *Journal of Economic Dynamics and Control*, Vol. 12:231–254.

[40] Johansen, S. (1991). Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. *Econometrica*, Vol. 59:1551–1580.

[41] Johansen, S. (1992a). Determination of Cointegration Rank in the Presence of a Linear Trend. *Oxford Bulletin of Economics and Statistics*, Vol. 54:383–397.

[42] Johansen, S. (1992b). Testing Weak Exogeneity and the Order of Cointegration in UK Money Demand Data. *Journal of Policy Modeling*, Vol. 14:313–334.

[43] Johansen, S. (1992c). Cointegration in Partial Systems and the Efficiency of Single-Equation Analysis. *Journal of Econometrics*, Vol. 52:389–402.

[44] Johansen, S. (1994). The Role of the Constant and Linear Terms in Cointegration Analysis of Nonstationary Variables. *Econometric Reviews*, Vol. 13:205–229.

[45] Johansen, S. and K. Juselius (1990). Maximum Likelihood Estimation and Inference on Cointegration - with Application to the Demand for Money. *Oxford Bulletin of Economics and Statistics*, Vol. 52:169–210.

[46] Juselius, K. (2006). *The Cointegrated VAR model: Methodology and Applications.* Oxford University Press, Oxford.

[47] Karfakis, C. I. and A. Parikh (1994). Exchange Rate Convergence and Market Efficiency. *Applied Financial Economics*, Vol. 4:93–98.

[48] Kwiatkowski, D.; P. C. B. Phillips; P. Schmidt and Y. Shin (1992). Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root : How sure are we that Economic Time Series have a Unit Root? *Journal of Econometrics*, Vol. 54:159–178.

[49] Lajaunie, J. P. and A. Naka (1992). Is the Tokyo Spot Foreign Exchange Market Consistent with the Efficient Market Hypothesis? *Review of Financial Economics*, Vol. 2:68–74.

[50] Lajaunie, J. P. and A. Naka (1997). Re-Examining Cointegration, Unit Roots and Efficiency in Foreign Exchange Rates. *Journal of Business Finance & Accounting*, Vol. 24:363–374.

[51] Lajaunie, J.P., B. L. McManis and A. Naka (1996). Further Evidence on Foreign Exchange Market Efficiency: An Application of Cointegration Tests. *The Financial Review*, Vol. 31:553–564.

[52] Lee, T.-H. and Y. Tse (1996). Cointegration Tests with Conditional Heteroskedasticity. *Journal of Econometrics*, Vol. 73:401–410.

[53] LeRoy, S. F. (1989). Efficient Capital Markets and Martingales. *Journal of Economic Literature*, Vol. 27:1583–1621.

[54] Levich, R. M. (1979). On the Efficency of Markets for Foreign Exchange. In : R. Dornbusch und J. A. Frenkel, editor, *International Economic Policy - Theory and Evidence*, Baltimore, London.

[55] Levich, R. M. (1985). Empirical Studies of Echange Rates: Price Behavior, Rate Determination and Market Efficiency. In : R.W. Jones und P.B. Kenen, editor, *Handbook of International Economics, Vol. II*, Amsterdam.

[56] Lütkepohl, H. (2005). *New Introduction to Multiple Time Series Analysis*. Springer-Verlag, Berlin.

[57] MacDonald, R. and M. P. Taylor (1989). Foreign Exchange Market Efficiency and Cointegration - Some Evidence from the Recent Float. *Economics Letters*, Vol. 29:63–68.

[58] MacKinnon, J. G. (1991). Critical Values for Cointegration Tests in Long-Run Economic Relationships. In : R. F. Engle und C. W. J. Granger, editor, *Readings in Cointegration*, New York.

[59] Masih, A. M.M. and R. Masih (1994). On the Robustness of Cointegration Tests of the Market Efficiency Hypothesis: Evidence from Six European Foreign Exchange Markets. *Economia internazionale: Journal of the Institute for International Economic*, Vol. 47:161–177.

[60] Meese, R. A. and K. J. Singleton (1982). On Unit Roots and the Empirical Modeling of Exchange Rates. *The Journal of Finance*, Vol. 37:1029–1035.

[61] Norrbin, S. C. (1996). Bivariate Cointegration among European Monetary System Exchange Rates. *Applied Economics*, Vol. 28:1505–1513.

[62] Osterwald-Lenum, M. (1992). A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics. *Oxford Bulletin of Economics and Statistics*, Vol. 54:461–471.

[63] Phengpis, C. (2006). Market Efficiency and Cointegration of Spot Exchange Rates during Periods of Economic Turmoil: Another Look at European and Asian Currency Crises. *Journal of Econonomics and Business*, Vol. 58:323–342.

[64] Rahbek, A.; E. Hansen and J. G. Dennis (2002). ARCH Innovations and their Impact on Cointegration Rank Testing. *Department of Theoretical Statistics, University of Copenhagen, Centre for Analytical Finance Working Paper*, WP No. 22.

[65] Rangvid, J. and C. Sorensen (2002). Convergence in the ERM and Declining Numbers of Common Stochastic Trend. *Journal of Emerging Market Finance*, Vol. 1:183–213.

[66] Rapp, T. A. and S. C. Sharma (1999). Exchange Rate Market Efficiency: Across and Within Countries. *Journal of Economics and Business*, Vol. 51:423–439.

[67] Reimers, H.-E. (1992). Comparisons of Tests for Multivariate Cointegration. *Statistical Papers*, Vol. 33:335–359.

[68] Richards, A. J. (1995). Comovements in National Stock Market Returns: Evidence of Predictability, but no Cointegration. *Journal of Monetary Economics*, Vol. 36:631–654.

[69] Schwert, G. W. (1989). Tests for unit roots: a Monte Carlo investigation. *Journal of Business and Economic Statistics*, Vol. 7:147–159.

[70] Sephton, P. S. and H. K. Larsen (1991). Tests of Exchange Market Efficiency: Fragile Evidence from Cointegration Tests. *Journal of International Money and Finance*, Vol. 10:561–570.

[71] Taylor, M.P. (1995). The Economics of Exchange Rates. *Journal of Economic Literature*, Vol. 33:13–47.

[72] Woo, K.-Y. (1999). Cointegration Analysis of the Intensity of the ERM Currencies under the European Monetary System. *Journal of International Financial Markets, Institutions and Money*, Vol. 9:393–405.

Bisher erschienene Diskussionspapiere

Nr. 68: Kühl, Michael: Cointegration in the Foreign Exchange Market and Market Efficiency since the Introduction of the Euro: Evidence based on bivariate Cointegration Analyses, Oktober 2007

Nr. 67: Hess, Sebastian, Cramon-Taubadel, Stephan von: Assessing General and Partial Equilibrium Simulations of Doha Round Outcomes using Meta-Analysis, August 2007

Nr. 66: Eckel, Carsten: International Trade and Retailing: Diversity versus Accessibility and the Creation of “Retail Deserts”, August 2007

Nr. 65: Stoschek, Barbara: The Political Economy of Enviromental Regulations and Industry Compensation, Juni 2007

Nr. 64: Martinez-Zarzoso, Inmaculada; Nowak-Lehmann D., Felicitas; Vollmer, Sebastian: The Log of Gravity Revisited, Juni 2007

Nr. 63: Gundel, Sebastian: Declining Export Prices due to Increased Competition from NIC – Evidence from Germany and the CEEC, April 2007

Nr. 62: Wilckens, Sebastian: Should WTO Dispute Settlement Be Subsidized?, April 2007

Nr. 61: Schöller, Deborah: Service Offshoring: A Challenge for Employment? Evidence from Germany, April 2007

Nr. 60: Janeba, Eckhard: Exports, Unemployment and the Welfare State, März 2007

Nr. 59: Lambsdoff, Johann Graf; Nell, Mathias: Fighting Corruption with Asymmetric Penalties and Leniency, Februar 2007

Nr. 58: Köller, Mareike: Unterschiedliche Direktinvestitionen in Irland – Eine theoriegestützte Analyse, August 2006

Nr. 57: Entorf, Horst; Lauk, Martina: Peer Effects, Social Multipliers and Migrants at School: An International Comparison, März 2007 (revidierte Fassung von Juli 2006)

Nr. 56: Görlich, Dennis; Trebesch, Christoph: Mass Migration and Seasonality Evidence on Moldova’s Labour Exodus, Mai 2006

Nr. 55: Brandmeier, Michael: Reasons for Real Appreciation in Central Europe, Mai 2006

Nr. 54: Martínez-Zarzoso, Inmaculada; Nowak-Lehmann D., Felicitas: Is Distance a Good Proxy for Transport Costs? The Case of Competing Transport Modes, Mai 2006

Nr. 53: Ahrens, Joachim; Ohr, Renate; Zeddies, Götz: Enhanced Cooperation in an Enlarged EU, April 2006

Nr. 52: Stöwhase, Sven: Discrete Investment and Tax Competition when Firms shift Profits, April 2006

Nr. 51: Pelzer, Gesa: Darstellung der Beschäftigungseffekte von Exporten anhand einer Input-Output-Analyse, April 2006

Nr. 50: Elschner, Christina; Schwager, Robert: A Simulation Method to Measure the Tax Burden on Highly Skilled Manpower, März 2006

Nr. 49: Gaertner, Wulf; Xu, Yongsheng: A New Measure of the Standard of Living Based on Functionings, Oktober 2005

Nr. 48: Rincke, Johannes; Schwager, Robert: Skills, Social Mobility, and the Support for the Welfare State, September 2005

Nr. 47: Bose, Niloy; Neumann, Rebecca: Explaining the Trend and the Diversity in the Evolution of the Stock Market, Juli 2005

Nr. 46: Kleinert, Jörn; Toubal, Farid: Gravity for FDI, Juni 2005

Nr. 45: Eckel, Carsten: International Trade, Flexible Manufacturing and Outsourcing, Mai 2005

Nr. 44: Hafner, Kurt A.: International Patent Pattern and Technology Diffusion, Mai 2005

Nr. 43: Nowak-Lehmann D., Felicitas; Herzer, Dierk; Martínez-Zarzoso, Inmaculada; Vollmer, Sebastian: Turkey and the Ankara Treaty of 1963: What can Trade Integration Do for Turkish Exports, Mai 2005

Nr. 42: Südekum, Jens: Does the Home Market Effect Arise in a Three-Country Model?, April 2005

Nr. 41: Carlberg, Michael: International Monetary Policy Coordination, April 2005

Nr. 40: Herzog, Bodo: Why do bigger countries have more problems with the Stability and Growth Pact?, April 2005

Nr. 39: Marouani, Mohamed A.: The Impact of the Mulitfiber Agreement Phaseout on Unemployment in Tunisia: a Prospective Dynamic Analysis, Januar 2005

Nr. 38: Bauer, Philipp; Riphahn, Regina T.: Heterogeneity in the Intergenerational Transmission of Educational Attainment: Evidence from Switzerland on Natives and Second Generation Immigrants, Januar 2005

Nr. 37: Büttner, Thiess: The Incentive Effect of Fiscal Equalization Transfers on Tax Policy, Januar 2005

Nr. 36: Feuerstein, Switgard; Grimm, Oliver: On the Credibility of Currency Boards, Oktober 2004

Nr. 35: Michaelis, Jochen; Minich, Heike: Inflationsdifferenzen im Euroraum – eine Bestandsaufnahme, Oktober 2004

Nr. 34: Neary, J. Peter: Cross-Border Mergers as Instruments of Comparative Advantage, Juli 2004

Nr. 33: Bjorvatn, Kjetil; Cappelen, Alexander W.: Globalisation, inequality and redistribution, Juli 2004

Nr. 32: Stremmel, Dennis: Geistige Eigentumsrechte im Welthandel: Stellt das TRIPs-Abkommen ein Protektionsinstrument der Industrieländer dar?, Juli 2004

Nr. 31: Hafner, Kurt: Industrial Agglomeration and Economic Development, Juni 2004

Nr. 30: Martinez-Zarzoso, Inmaculada; Nowak-Lehmann D., Felicitas: MERCOSUR-European Union Trade: How Important is EU Trade Liberalisation for MERCOSUR's Exports?, Juni 2004

Nr. 29: Birk, Angela; Michaelis, Jochen: Employment- and Growth Effects of Tax Reforms, Juni 2004

Nr. 28: Broll, Udo; Hansen, Sabine: Labour Demand and Exchange Rate Volatility, Juni 2004

Nr. 27: Bofinger, Peter; Mayer, Eric: Monetary and Fiscal Policy Interaction in the Euro Area with different assumptions on the Phillips curve, Juni 2004

Nr. 26: Torlak, Elvisa: Foreign Direct Investment, Technology Transfer and Productivity Growth in Transition Countries, Juni 2004

Nr. 25: Lorz, Oliver; Willmann, Gerald: On the Endogenous Allocation of Decision Powers in Federal Structures, Juni 2004

Nr. 24: Felbermayr, Gabriel J.: Specialization on a Technologically Stagnant Sector Need Not Be Bad for Growth, Juni 2004

Nr. 23: Carlberg, Michael: Monetary and Fiscal Policy Interactions in the Euro Area, Juni 2004

Nr. 22: Stähler, Frank: Market Entry and Foreign Direct Investment, Januar 2004

Nr. 21: Bester, Helmut; Konrad, Kai A.: Easy Targets and the Timing of Conflict, Dezember 2003

Nr. 20: Eckel, Carsten: Does globalization lead to specialization, November 2003

Nr. 19: Ohr, Renate; Schmidt, André: Der Stabilitäts- und Wachstumspakt im Zielkonflikt zwischen fiskalischer Flexibilität und Glaubwürdigkeit: Ein Reform-ansatz unter Berücksichtigung konstitutionen- und institutionenökonomischer Aspekte, August 2003

Nr. 18: Ruehmann, Peter: Der deutsche Arbeitsmarkt: Fehlentwicklungen, Ursachen und Reformansätze, August 2003

Nr. 17: Suedekum, Jens: Subsidizing Education in the Economic Periphery: Another Pitfall of Regional Policies?, Januar 2003

Nr. 16: Graf Lambsdorff, Johann; Schinke, Michael: Non-Benevolent Central Banks, Dezember 2002

Nr. 15: Ziltener, Patrick: Wirtschaftliche Effekte des EU-Binnenmarktprogramms, November 2002

Nr. 14: Haufler, Andreas; Wooton, Ian: Regional Tax Coordination and Foreign Direct Investment, November 2001

Nr. 13: Schmidt, André: Non-Competition Factors in the European Competition Policy: The Necessity of Institutional Reforms, August 2001

Nr. 12: Lewis, Mervyn K.: Risk Management in Public Private Partnerships, Juni 2001

Nr. 11: Haaland, Jan I.; Wooton, Ian: Multinational Firms: Easy Come, Easy Go?, Mai 2001

Nr. 10: Wilkens, Ingrid: Flexibilisierung der Arbeit in den Niederlanden: Die Entwicklung atypischer Beschäftigung unter Berücksichtigung der Frauenerwerbstätigkeit, Januar 2001

Nr. 9: Graf Lambsdorff, Johann: How Corruption in Government Affects Public Welfare – A Review of Theories, Januar 2001

Nr. 8: Angermüller, Niels-Olaf: Währungskrisenmodelle aus neuerer Sicht, Oktober 2000

Nr. 7: Nowak-Lehmann, Felicitas: Was there Endogenous Growth in Chile (1960-1998)? A Test of the AK model, Oktober 2000

Nr. 6: Lunn, John; Steen, Todd P.: The Heterogeneity of Self-Employment: The Example of Asians in the United States, Juli 2000

Nr. 5: Güßefeldt, Jörg; Streit, Clemens: Disparitäten regionalwirtschaftlicher Entwicklung in der EU, Mai 2000

Nr. 4: Haufler, Andreas: Corporate Taxation, Profit Shifting, and the Efficiency of Public Input Provision, 1999

Nr. 3: Rühmann, Peter: European Monetary Union and National Labour Markets, September 1999

Nr. 2: Jarchow, Hans-Joachim: Eine offene Volkswirtschaft unter Berücksichtigung des Aktienmarktes, 1999

Nr. 1: Padoa-Schioppa, Tommaso: Reflections on the Globalization and the Europeanization of the Economy, Juni 1999

Alle bisher erschienenen Diskussionspapiere zum Download finden Sie im Internet unter: http://www.uni-goettingen.de/de/60920.html.