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Keskusteluaiheita - Discussion papers

No. 574

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THE FORWARD EXCHANGE RATE AS A PREDICTOR OF THE SPOT EXCHANGE RATE

An empirical study

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ISSN 0781-6847

03.12.1996

HAILI, Heidi, **THE FORWARD EXCHANGE RATE AS A PREDICTOR OF THE SPOT EXCHANGE RATE, An empirical study**. Helsinki: ETLA, Elinkeinoelämän Tutkimuslaitos, The Research Institute of the Finnish Economy, 1996, 88 p. (Keskusteluaiheita, Discussion Papers, ISSN 0781-6847; no. 574).

ABSTRACT: In this empirical study we wish to assess whether the forward exchange rates contain information in respect to the spot exchange rates. The starting point of the study is the unbiased efficient expectations hypothesis (UEE) which states, with the assumptions of interest rate parity and rational expectations, that the n-period forward rate at time t is the unbiased predictor of the spot rate at time t+n, given the information set and possible risk premium known at time t. The data represents the one, three and six month FIM/USD and FIM/DEM forward exchange rates with the corresponding spot rates. In order to study the link between the forward and spot rates, we use the maximum likelihood cointegration analysis by Johansen (1988a). The analysis allows the non-stationary nature of the rates and establishes a long run equilibrium relationship between them. The main finding of the study is the establishment of the above mentioned equilibrium relationships between the forward and spot rates, indicating that the forward rates do contain information pertinent to the determination of the spot rates. We cannot, however, confirm the existence of an unbiased relationship between the spot rate and the three forward rates.

KEY WORDS: Forward exchange rate, spot exchange rate, efficient markets, non-stationarity, cointegration

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TIIVISTELMÄ: Tässä empiirisessä tutkimuksessa tutkitaan sisältävätkö termiinivaluuttakurssit informaatiota spot-valuuttakurssien määräytymisen suhteen. Lähtökohtana tutkimukselle on harhaton tehokkaiden odotusten hypoteesi (the unbiased efficient expectations hypothesis, UEE), jonka mukaan, olettaen korkopariteetin ja rationaalisten odotusten olemassaolon, n-periodin termiinivaluuttakurssi ajankohtana t on ajankohdan t+n spot-valuuttakurssin harhaton estimaatti, ottaen huomioon markkinainformaation ja mahdollisen riskipreemion hetkellä t. Aineistona käytetään päivittäisiä yhden-, kolmen- ja kuuden kuukauden FIM/USD- ja FIM/DEM-termiinivaluuttakursseja sekä vastaavia spot-valuuttakursseja. Termiini- ja spot-valuuttakurssien yhteyttä tutkitaan Johansenin (1988a) suurimman uskottavuuden yhteisintegroituvuusmenetelmällä. Menetelmä ottaa huomioon valuuttakurssien epästationaarisen luonteen ja menetelmän avulla voidaan muodostaa kurssien välille pitkän aikavälin tasapainoyhtälö. Tutkimuksessa voidaan vahvistaa tasapainoyhtälöiden olemassaolo termiini- ja spot-valuuttakurssien välille. Tämä tarkoittaa sitä, että termiinikurssit sisältävät spot-kurssien määrätymiseen vaikuttavaa informaatiota. Tutkimuksessa ei kuitenkaan hyväksytä hypoteesia, jonka mukaan kurssien välillä vallitsee harhaton ennusteyhteys.

ASIASANAT: Termiinivaluuttakurssi, spot-valuuttakurssi, tehokkaat markkinat, epästationaarisuus, yhteisintegroituvuus

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Yhteenveto - Finnish Summary

Tutkimuksen tarkoituksena on selvittää, sisältävätkö termiinivaluuttakurssit informaatiota tulevaisuuden spot-valuuttakursseista. Tämä on yksi kansainvälisen rahoitusteorian klassisista kysymyksistä. Jos markkinat toimivat tehokkaasti, eli kaikki markkinoillatoimijoiden käytettävissä oleva informaatio vaikuttaa valuuttakurssien arvon määräytymiseen, termiinivaluuttakurssi on harhaton estimaatti spot-valuuttakurssille.

Saatujen tulosten mukaan on selvää, että termiinivaluuttakurssin arvo vaikuttaa tulevaisuudessa toteutuvan vastaavan spot-valuuttakurssin arvoon. Kuitenkaan kaikissa tapauksissa ei voida todentaa, että termiinivaluuttakurssi ja tulevaisuuden spot-valuuttakurssi olisivat yhteneväiset. Täten Suomen valuuttakurssimarkkinoiden tehokkuutta on ehkä syytä epäillä. Yhteneväisyys ilmenee vain lyhyimmällä, yhden kuukauden, ennusteperiodilla, jolloin tämä voidaan selittää sillä, että uutta, valuuttakurssien arvoon vaikuttavaa informaatiota, joka ei vielä sisälly termiinivaluuttakurssin arvoon, ei enää ehdi syntyä niin paljon kuin esimerkiksi kuuden kuukauden ennusteperiodilla.

Aineistona tutkimuksessa käytettiin päivittäisiä Saksan markan ja Yhdysvaltain dollarin valuuttakursseja Suomen markkaa vastaan. Ennustaviksi termiinikursseiksi valittiin yhden-, kolmen- ja kuuden kuukauden termiinivaluuttakurssit. Koska pisin ennusteperiodi on kuusi kuukautta, tarkastelujakso alkaa kuusi kuukautta siitä hetkestä, jolloin markka päästettiin kellumaan, 8.9.1992. Tarkasteluajanjakso on täten 8.3.1993 - 29.4.1994. Tutkimuksessa haluttiin käyttää nimenomaan tarkasteluperiodia, jolloin markka ei ole kytkettynä mihinkään valuuttakurssijärjestelmään, vaan kelluu vapaasti ja siten heijastaa arvollaan vain markkainformaatiota. Tutkimuksessa käytetään kahden eri valuutan Suomen markan kurssia tutkimustulosten luotettavuuden lisäämiseksi.

Lähtökohtana käytetään em. tehokkaiden markkinoiden hypoteesin mukaista ns. harhattomien odotusten hypoteesia, jolla tarkoitetaan, että hetkellä t määräytyvä termiinivaluuttakurssi periodille n on hetken t + n spot-valuuttakurssin harhaton estimaatti siten, että ne ovat yhteneväiset. Tämä hypoteesi perustuu myös ns. korkopariteetin ja rationaalisten odotusten toteutumiselle. Korkopariteetilla tarkoitetaan, että koti- ja ulkomaisten korkojen tulee olla tasapainossa siten, että niiden ero kompensoituu ko. valuuttojen arvojen erolla. Rationaalinen odotus taas jonkin tapahtuman toteutumisesta perustuu ko. tapahtuman todella aikaansaavaan prosessiin, eikä mihin tahansa tapahtumaprosessiin.

Tutkimuksen empiirisessä mallintamisessa käytetään ns. Johansenin (1988a) suurimman uskottavuuden yhteisintegroituvuusmenetelmää, jolla pystytään muodostamaan pitkän aikavälin tasapainoyhtälö termiinivaluuttakurssien ja vastaavien spot-valuuttakurssien välille. Menetelmä myös mahdollistaa erilaisten rakenteellisten hypoteesien testaamisen aineistolla. Juuri tällä tavoin pystyttiin päättelemään, että termiinikurssit eivät ole spot-valuuttakurssien harhattomia estimaatteja.

1. INTRODUCTION

The problem this study will look into is the following: how well is it possible to forecast the Finnish markka daily spot exchange rate from the term structure of forward exchange rates? The link between the two exchange rates has not been studied with the Finnish data before. One reason is, that the Finnish markka has always been pegged to either another currency or a currency basket and floated freely[1] only since 8 September, 1992. In this study we will use the exchange rate of the Finnish markka against the German Mark and the US Dollar during period 8.3.1993 - 29.4.1994.[2]

The exchange rate policy of Finland after World War II was based on a fixed exchange rate. Several devaluations and revaluations of the markka have taken place, however. After the war the external value of the markka was fixed within the Bretton Woods system. When the Bretton Woods system broke up, the Finnish markka was pegged to a trade-weighted currency basket from 1 November, 1977 until 7 June, 1991. A regime change was made and the markka was pegged to the European Currency Unit, the ECU, from 7 June, 1991 until 7 September, 1992, when it was let float. During the ECU connection, the markka was devalued on 15 November, 1991 by 12.3%. The floating of the markka may last for an indefinite period on the basis of the amended Currency Act. (Bank of Finland Bulletin, 1995, 69:10, p.21) The period ended on 14 October, 1996, when the markka joined the Exchange Rate Mechanism (ERM).

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[1]The floating of the Finnish markka is best described by the term managed or "dirty" float because the Bank of Finland has admitted it stabilises any big movements in the exchange rate.
[2]The period start date is exactly six months after the floating date, because the longest horizon explanatory variable to the spot rate is the six-month forward exchange rate. The period end date was chosen randomly. The data are shown graphically in Appendices 2 and 3.

The problem of concordance of the spot and forward rates is not new. In the literature of international finance it has been considered from different aspects during the past 20 years with controversial results. The problem itself can be expressed as a combination of covered and uncovered interest rate parities and the efficient market hypothesis (EMH), and it can be called simple[3] efficiency hypothesis, unbiasedness hypothesis or the newest verbal description: unbiased efficient expectations (UEE) hypothesis[4] (expressed in natural logarithms):

$$(1.1) \qquad f_{n,t} = E(s_{t+n}|I_t)$$

Clarida & Taylor (1993, 1) formulate the hypothesis (1.1) as 'the equilibrium forward exchange rate established at time t for delivery of foreign exchange at date t+n . . . should be the best available predictor of the level the spot exchange rate realised at date t+n'. I_t in equation (1.1) is the information set known by the market at time t.

First the problem was studied by simply regressing the forward rate and information variables on the corresponding spot rate (see e.g. Frenkel 1981 and Taylor 1995). It was soon noted, however, that modelling non-stationary variables, like most macroeconomic data including the exchange rates, by ordinary least squares -method leads to spurious regression with dubious results. As a solution, researchers run regressions of the forward premium on the change in the spot exchange rate at time t. This method still confuses stationary and non-stationary variables. the real solution to the problem is the development of unit root econometrics and the concept of cointegration (see e.g. Engle & Granger 1987, Johansen 1988a, Johansen & Juselius 1990). The cointegration

[3]Simple in the sense that it allows for no risk premium (MacDonald and Taylor 1991, 26-27).
[4]Following Baillie & MacMahon (1989, 162).

method is not only a means of modelling non-stationary data, but it also provides a new perspective to conventional data modelling: it enables distinguishing long run relations from short run dynamic adjustment. This study is based on cointegration methodology. The methodology enables us to define a long run equilibrium relationship between the spot and forward exchange rates. The purpose of this study is not to maintain that the UEE hypothesis is true; we rather wish to find evidence whether the forward rates of different maturities contain information pertinent to the determination of the spot rate of the corresponding time horizon. Our study follows in a broad manner that of Clarida and Taylor (1993). Major differences include our emphasis on the model misspecification tests, which are ignored by Clarida and Taylor, and the parameter restriction tests on the estimated cointegration coefficients obtained by the Johansen maximum likelihood procedure for determining the cointegrating vectors. We do this in order to obtain more information about the relationship between the rates. In addition, it is interesting to conduct the study with several forward rates; most studies concentrate on one forward rate time horizon as a predictor for the particular spot rate. (e.g. Baillie & Bollerslev 1989 and Hakkio & Rush 1989)

After defining the theoretical concepts vital to the study in section one, we survey the earlier research of the subject in section two, which is followed by an introduction to the econometric methodology used in the study in section four. Our analysis of cointegration is started with unit root tests on the variables in section five, which concentrates on data and analysis on the whole. Then we conduct the cointegration tests and test the relevant restrictions on the estimated parameters. The study is concluded by section six.

2. TERMS AND CONCEPTS

The spot exchange rate, denoted as S, is the price of foreign currency in the domestic currency.

e.g. S = FIM 4.2 / USD 1

This notation requires that a devaluation or a depreciation of S against the foreign currency is shown as an increase in its numerical value. On the other hand a revaluation or an appreciation of S against the foreign currency implies a decrease in the numerical value of S. The notation S_t simply stands for the spot exchange rate at time t for delivery at time t, that is, on the spot.

The forward exchange rate, $F_{n,t}$, is the price of foreign currency in terms of domestic currency at time t for delivery of foreign exchange at time t+n. When the value of a spot exchange rate in a time, S_t, exceeds the value of a forward rate, $F_{n,t}$, the domestic currency is said to be selling at a premium and the difference between the two rates, $(F_{n,t} - S_t)$, is called *a forward premium*. The negative of the premium is a discount, then $(F_{n,t} - S_t) > 0$ and the currency is selling at a discount. In this paper we are only using the term forward premium when referring to the difference between forward and spot exchange rates.

2.1. Efficient Markets Hypothesis

The efficient markets hypothesis (EMH) was originally formulated for asset markets, but it is well applicable to foreign exchange markets since foreign exchange can be seen as a form of an asset. There are some problems associated with the hypothesis, however, which will be discussed later.

The concept of efficient markets was first introduced by Eugene F. Fama in his 1965 article. According to the article, the efficient market consists of a

> large number of rational profit maximisers actively competing with each other to predict future market values of individual securities and where important current information is almost freely available to all participants (Fama 1965, 34-105).

This means that the prices reflect all the relevant and available information and no participant in the market is able to make ex ante unexploited profits based on the market prices.

Fama also defines three levels of market efficiency based on how much information is understood to underlie the notion 'all the relevant and available information'. These three forms of market efficiency mean that no investor can make excess profits from trading rules based on the different types of information. The efficiency forms are:

i) *the weak form of efficiency*: prices reflect all information contained in the past prices.
ii) *the semistrong form of efficiency*: prices reflect not only past prices but also all other published information.

iii) *the strong form of efficiency*: prices reflect not only public information but in addition all information that can be acquired and that can have an effect on the price of an asset.

(Copeland & Weston 1988, 322 and Tucker et al. 1991, 46-48 and Brealey & Myers 1991, 295.) The strong form of efficiency does not (necessarily) exist in the market for foreign exchange, because for example the central banks can unexpectedly intervene in the market, which is hardly of public knowledge beforehand and thus cannot be discounted in the exchange rates. It should be noted, however, than in the case of central banks, the interventions can follow a certain pattern according to which the market participants may have formed intervention rules, which would make the interventions both expected and maybe unnecessary.

The above division, however, has received criticism and not least from the direction of its creator Fama himself. Not long ago (Fama 1991, 1575-1671) he modified the three forms of efficient markets to more testable patterns. The weak form tests should be tests for *return predictability* with past returns, volatility and asset specific characteristics as explanatory variables. Semistrong and strong forms of market efficiency should be replaced with *event studies* and tests for *private information*, respectively. These new definitions of market efficiency give more practical means for testing. But Fama (1991, 1575-1576) stresses the importance of understanding that testing for EMH involves testing the joint hypothesis of the degree of how well information is reflected in prices and how well some model of equilibrium works as a transmitter.

Given the definitions above, the efficient market hypothesis could actually be referred to as an information efficient market, because 'all relevant information' is immediately transferred to prices. The question is, however: how is the information transferred to prices, how are the expectations formed and what is the role of risk to an investor? Answers to these questions are discussed below.

2.2. Rational Expectations

Since rational expectations (RE) are vital to the efficient market hypothesis, we will try to explain their meaning in expectations formation.

To start with, rational expectations does not mean that the decision maker always makes the right decision. Right decisions are as stochastic as prediction errors. RE does not even require that every participant in the market has the same expectations. RE is a methodological, model specific principle. (Vilmunen 1994) But usually the rational expectation on an event is a more accurate prediction than a prediction based on another basis, because RE is based on the process actually determining the process. (Attfield et al. 1991, 24)

In the context of exchange rates, and more specifically, the spot exchange rate, RE can be stated as:

(2.1) $$\Delta s_t = \Delta s_t^e + \eta_t$$

where s stands for the spot exchange rate expressed in natural logarithms, $\Delta s_t = s_t - s_{t-1}$, $\Delta s_t^e = s_t^e - s_{t-1}$ and η_t is a random forecast error, orthogonal to the information set. The expectations process is formed in the following way:

$$\Delta s_t^e = E(\Delta s_t | I_{t-1})$$

where the notation E(.) denotes the mathematical expectation conditioned on the information set, I_{t-1}, available to economic agents at time t. (MacDonald and Taylor 1991, 26)

If participants in the market are risk neutral, that is they regard domestic and foreign assets as identical in terms of maturity and risk, profits can be made by taking open forward positions if the forward rate at time t for delivery n periods later differs from the expected spot rate at time t+n. This forces the rates into equality:

(2.2) $$f_{n,t} = E(s_{t+n} | I_t)$$

(MacDonald and Taylor 1991, 26-27). This is called the simple efficiency hypothesis, because it does not take into account a possible risk premium. (Clarida & Taylor 1993, 12) or the unbiased efficient expectations (UEE) hypothesis for much the same reason plus the RE implication of efficient expectations (Baillie & MacMahon 1989, 163-164).

If market participants are risk averse, a risk premium, λ_t, is added to equation (2.2) so as to compensate for taking open forward positions:

(2.3) $$f_{n,t} = E(s_{t+n} | I_t) + \lambda_t$$

The strict form of market efficiency, (2.2), which allows for no risk premium, has been examined widely. As a consequence, researchers have come to draw the conclusion that a risk premium is valid, because they have not been otherwise able to explain the failure of their tests from the EMH. It should be noted however, that the EMH is a joint hypothesis of rational expectations and the chosen risk behaviour of the market participants. This means that the rejection EMH can be caused by either of these assumptions. As to the nature of the risk premium, there is evidence of a time varying risk premium, which, according to MacDonald and Taylor (1991, 30-31), is not a plausible explanation so far. They suggest that the failure of EMH could be better explained by a failure in the rational expectations. Actually Froot and Frankel (1989, quoted in Clarida & Taylor 1993, 8) have shown by using survey data of exchange rates that the bias between forward and spot exchange rates, like λ_t in (2.3), is not entirely due to risk. They draw the conclusion that the bias is caused by systematic expectational errors.

2.3. Shortcomings of the Efficient Markets Hypothesis

The time series of the spot and forward exchange rates in this study are expressed in natural logarithms, that is $f_t = \ln F_t$ and $s_t = \ln S_t$. The reason is the so called 'Siegel's paradox'. Siegel (1972) noted that the unbiased efficient expectations (UEE) hypothesis[5]

$$(2.4) \qquad F_t = E(S_{t+1})$$

[5] In 1972 the hypothesis was known as the unbiasedness hypothesis. Later it was renamed the speculative efficiency hypothesis and efficient markets hypothesis (Baillie & McMahon 1989, 163-164). Today we prefer the term unbiased efficient expectations hypothesis.

does not hold in both directions when the same exchange rate is expressed from the point of view of the foreign country as

$$(2.5) \qquad \left[\frac{1}{F_t}\right] = E\left[\frac{1}{S_{t+1}}\right]$$

The reason for this is the Jensen's inequality, which implies for convex functions that

$$(2.6) \qquad E\left(\frac{1}{x}\right) > \left(\frac{1}{E(x)}\right)$$

The problem is solved by taking natural logarithms of the time series.[6] (Beenstock 1985)

Another source of criticism towards the EMH arises from its assumption of freely available information and a similar model for every agent in the market which they use to form their price predictions by rational expectations. Grossman and Stiglitz (1980, quoted in Baillie & MacMahon 1989, 50-51) have developed an 'impossibility theorem', which deals with the relation between the transmission of information and the informational efficiency of markets. According to the theorem, it would not be in the interest of any participant in the market to acquire new information if it was costless because, since it is already reflected in the prices, there would be no reward for

[6]The simple EMH in logarithmic form is f_t=E(s_{t+1}) which from the point of view of the foreign country translates into 1-f_t=E(1-s_{t+1}) and further to f_t=E(s_{t+1}). This means that the logarithmic expression for the exchange rate is independent of the way the rates are quoted.

acquiring information. If information is costly, the market could not be efficient and prices could not reflect all the information that is available because then those who extract the information remain without profit. Otherwise there would exist a so called free rider problem. In equilibrium the marginal revenue from new information should equal the marginal cost of acquiring the information.

However, we must conclude the criticism of EMH by noting that several[7] researchers still recently have comfortably used the efficient market hypothesis as a basis for their research. Therefore we see no hindrance in doing so.

2.4. Interest Rate Parity Theorem

The interest rate parities (IRP) provide a model for exchange rate determination, a link between domestic and foreign interest rates and spot and forward markets. (See e.g. Copeland 1989, 86-91 or Baillie & MacMahon 1989, 150-151.)

The uncovered interest rate parity (UIRP) condition assumes that the market participants, investors, are risk neutral in relation to domestic and foreign assets. The assets must have the same risk characteristics and market imperfections, for instance, because transaction costs are not allowed. The expected returns on both the assets are

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[7]To name a few: Taylor (1988), Hakkio & Rush (1989), MacDonald & Taylor (1989), Lajaunie & Naka (1992), Clarida & Taylor (1993), Sulamaa (1993).

equal, when

$$(2.7) \qquad \frac{E(S_{t+n}) - S_t}{S_t} = \frac{i_t - i_t^*}{1 + i_t^*}$$

Equation (2.7) holds that the expected change in the spot exchange rate at time t, given the information known at that time (the left hand side of the equation) must equal the difference between domestic and foreign interest rates (i and i*, respectively) for the relevant maturity (n), since $(1 + i_t^*) \approx 1$. It is thus imperative that the domestic interest rate equals the sum of the foreign interest rate and the expected rate of depreciation of the domestic currency.

Equation (2.7) is called the uncovered interest rate parity (UIRP) condition because the risk on the expected spot exchange rate is not covered. In the covered interest rate parity (CIRP) the risk is covered in the forward market by replacing the expected spot rate at time (t+n) by the forward rate, which is known at time t for delivery at time t+n. Here we do not have to make any assumptions on the risk behaviour of the investors.

$$(2.8) \qquad \frac{F_{n,t} - S_t}{S_t} = \frac{i_t - i_t^*}{1 + i_t^*}$$

The CIRP is an arbitrage condition, which means that arbitrage yields no profit.

Combining the uncovered and covered interest rate parities gives us a statement for the unbiased efficient expectations (UEE) hypothesis which was discussed in section 2.2:

$$(2.9) \qquad F_{n,t} = E(S_{t+n})$$

The above hypothesis states that the equilibrium forward exchange rate established now for delivery of foreign exchange n periods later is the best available estimator of the level of the spot exchange rate realised n periods later.

When (2.9) is combined with the RE statement of spot and forward exchange rates (2.2) or (2.3), the result is equation (2.10) below

$$(2.10) \qquad F_{n,t} = E(S_{t+n}|I_t) + \lambda_t$$

which is an informationally efficient spot rate with ($\lambda_t \neq 0$) or without ($\lambda_t = 0$) a risk premium.

3. SURVEY ON EARLIER RESEARCH

3.1. The difference between stationary and non-stationary variables

The question of how good a predictor of the spot exchange rate the forward rate really is, has challenged many researchers for the last decades. In 1981 Frenkel tested the hypothesis presented by (1.1) repeated below, by considering log-linear regressions like equation (3.1)

$$(1.1) \qquad E(s_{t+n}|I_t) = f_{n,t}$$

$$(3.1) \qquad s_{t+n} = \alpha + \beta f_{n,t} + \gamma z_t + \varepsilon_{t+n}$$

where the lower case letters denote variables in natural logarithms (applies to the whole study) and z_t is a vector of information variables known at time t. The null-hypothesis that the forward exchange rate is an unbiased estimator of the spot rate, i.e. that β=1 and γ=0, could not be rejected. This supports the UEE (1.1).

It was soon discovered, though, that there is a theoretical problem in regressions such as (3.1). The problem is that, like most macroeconomic time series, the spot and forward exchange rates are non-stationary[8] variables and treating them as stationary as in (3.1) leads to so called spurious regression with results indicating a meaningful causal relation between the variables when in reality there is only a coexistent correlation (Harris 1995, 14). A new technique was thus developed to find out the impact of the forward exchange rate on the spot exchange rate: a regression of the change in the spot rate on the forward premium and other variables. For instance Bilson (1981) and Fama (1984) run regressions of the form:

$$(3.2) \qquad s_{t+n} - s_t = \alpha + \beta(f_{n,t} - s_t) + \varepsilon_{t+n}$$

[8]A stationary process has a constant mean and variance (see section 4.1.).

Where $(f_{n,t} - s_t)$ is the forward premium (see section 2). The null-hypothesis was that β=1, but it was rejected and seen that, on the contrary, β=-0.88[9]. This result suggests in distinct contrast to the result of the earlier model (3.1) that the forward premium actually mispredicts the direction of the following change in the spot rate. For example when the domestic currency is selling at a discount, i.e. the difference between forward and spot exchange rates is positive, the interest rate parity theorem (see section 2.4.) suggests that the interest rate differential as compared to the foreign currency will grow. These results thus not only imply that the forward rate is not the optimal predictor of the spot exchange rate but also mean a rejection of the UEE and the interest rate parity theorem.

3.2. Cointegration in the study of UEE

Baillie and Selover (1987) were among the first to apply the concept of cointegration into the research on exchange rates. They analysed the monetary model of exchange rate determination[10] along the lines of Engle and Granger (1987) and were not able to find cointegrating relationships between the determinants of the model, which would have been a sign of long run relationships. However, it is interesting to note that they

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[9] β=-0.88 is the average of estimated values of β on regressions such as (4.2) in 75 published estimates (Clarida & Taylor 1993, 2).

[10] The pure monetary model:

$$s_t = \beta_1\left(m_t - m_t^*\right) + \beta_2\left(y_t - y_t^*\right) + \beta_3\left(r_t - r_t^*\right) + \beta_4 E_t\left(p_{t+1} - p_{t+1}^*\right) + u_t$$

where s is the natural logarithm of the nominal exchange rate, m and y natural logarithms of domestic money supply and real output respectively, r is the domestic short term interest rate, $E_t p_{t+1}$ is the expected domestic rate of inflation. The asterisks represent foreign respective variables and u_t is a stochastic disturbance term. (Baillie & Selover 1987, 44) were able to establish that the exchange rates in their data[11] were all integrated of order one, ie. non-stationary.

Baillie continued the cointegration analysis with Bollerslev (1989) in the context of unbiasedness hypothesis, i.e. the forward rate is an unbiased predictor of the spot exchange rate. First they again, using techniques introduced by Phillips and Perron (1988) found the time series of spot and forward rates for the seven countries under examination to be integrated of order one[12] and moved on to test whether a long run equilibrium relationship could be found between the seven sets of spot and forward exchange rates. They did this following the Engle-Granger two-step procedure (Engle & Granger 1987, see section 4.2. for an introduction) by first estimating equation (3.3) by ordinary least squares (OLS)

$$(3.3) \qquad y_{t+k} = s_{t+k} - a - bf_t$$

where y_{t+k} is a transitory equilibrium error, s_{t+k} is the spot rate and f_t the corresponding forward rate, to obtain

$$(3.4) \qquad s_{t+22} = \hat{a} + \hat{b}f_t + \hat{y}_{t+22}$$

where $\hat{y}_{t+22}$ denotes the OLS residual and k is chosen to be 22, because it indicates the number of working days in the context of daily spot rates and 30-day forward rates[13]

[11]The data consists of 130 end of month observations of spot exchange rates USD/GBP, USD/YEN, USD/DEM, USD/CAD, USD/FRF during period March 1973 until December 1983.
[12]They studied daily spot and one-month forward exchange rates for 1980-1985 in seven countries, namely: UK, West Germany, France, Italy, Switzerland, Japan, and Canada.
[13]30-(4*2)=22, where (4*2) are the weekend days in a month.

and is therefore the time period over which the expectation typically takes place. Baillie and Bollerslev then examine whether the OLS residual, $\hat{y}_t$, contains a unit root or not. They are not able to find one, and thus interpret the spot and forward rates to be cointegrated with a stationary equilibrium error, i.e. the forward premium is stationary. In order for the forward rate to be an unbiased predictor of the spot rate, a and b in (3.1) must equal zero and one, respectively; (a, b) = (0, 1). Conducted t-tests give results supportive of the hypothesis. After having assessed the cointegrating nature of the rates, Baillie and Bollerslev turn to another problem of market efficiency, which can very well be solved by using a similar technique: if exchange rates are determined in efficient markets, the price of one currency cannot affect the price of another, i.e. there should be no cointegrating relationship between the prices of two or more currencies. The theoretical basis for this is the Granger Representation Theorem (Granger 1986 and Engle and Granger 1987, introduced in section 4.1.) which implies an error-correction representation of the variables of the following form where the cointegrating relationship is the error correction term x_{t-1}:

$$\Delta s_t^a = -\rho_1 x_{t-1} + \sum_{i=1}^{n} \alpha_i \Delta s_{t-i}^a + \sum_{i=1}^{n} \beta_i \Delta s_{t-i}^b + \varepsilon_{1,t}$$

(3.5)

$$\Delta s_t^b = -\rho_2 x_{t-1} + \sum_{i=1}^{n} \phi_i \Delta s_{t-i}^a + \sum_{i=1}^{n} \lambda_i \Delta s_{t-i}^b + \varepsilon_{2,t}$$

where s^a and s^b are the two spot rates, $x_{t-1} = s_{t-1}^a - \beta s_{t-1}^b$ is the error-correction term, and $\rho_1 + \rho_2 \neq 0$. (MacDonald & Taylor 1989, 64) It is easy to see from (3.5) what was already explained above: if the error-correction equations between the variables exist, one can be used to forecast another.

The assumption that any exchange rate markets are not interdependent naturally presumes that the currencies are different assets (Hakkio & Rush 1989,78). If countries are similar in terms of production technologies and implicitly link their economic policies or alternatively fix their exchange rates explicitly, their currencies cannot be viewed as different assets. As to cointegration, currencies that are not regarded as different assets would then be cointegrated. A multivariate test due to Johansen (as in Baillie & Bollerslev 1989, 174) is conducted for the seven daily spot and forward rates, respectively and the result implies that 'the seven exchange rates are tied together in one long run relationship' (Baillie & Bollerslev 1989, 176). This means that the disequilibrium error from the relationship is an important factor in the exchange rate change in the next period. The weak form of market efficiency [14] is thus violated.

A study in a similar framework concerning the efficiency of a particular currency market in relation to other currency markets[15] is also conducted by MacDonald and Taylor (1989) and Coleman (1990)[16]. They all start by first determining the spot and forward exchange rates non-stationary and then turn to cointegration tests using the Engle-Granger two-step procedure and conclude with a contradicting result as

[14]The weak form of market efficiency: $s_t = a + bs_{t-1} + \varepsilon_t$

[15]For instance: is there a connection between the Finnish and the Swedish currency markets?

[16]MacDonald and Taylor (1989) use the spot exchange rate data for 10 currencies (AUD, BEF, DKK, FRF, DEM, ITL, NLG, CAD, YEN, and GBP) from January 1973 until December 1985, collected monthly, i.e. about 150 observations on each exchange rate. As to Coleman (1990), he has collected daily spot exchange rate data for 18 currencies (ATS, BEF, GBP, CAD, DKK, NLG, FIM, FRF, HKD, ITL, JPY, NZD, NOK, SGD, ESP, SFR, SEK, and DEM) from June 1973 or January 1976 until presumably at least December 1985 (not mentioned in the paper).

compared to Baillie and Bollerslev (1989) above: there is no evidence of the interdependence of the markets.

An extensive paper on foreign exchange market efficiency and cointegration was prepared by Hakkio and Rush in 1989. Their definition of the efficient foreign exchange market assumes that 'agents are risk neutral and use all available information rationally so that the forward exchange rate is an unbiased predictor of the future spot rate' (Hakkio & Rush 1989, 75). They test the market efficiency in five different tests based on unit root tests introduced by Engle and Granger (1987) and the Engle-Granger two-step procedure for determining cointegration. First, they too, establish the non-stationarity of the spot and forward rates[17]. Second, they examine the German and British spot and forward markets, respectively, based on the same kind of reasoning as Baillie and Bollerslev (1989), MacDonald and Taylor (1989), and Coleman (1990) above. In contrast to Baillie and Bollerslev (1989) and in common with MacDonald and Taylor (1989), and Coleman (1990), Hakkio and Rush (1989) find the markets efficient by not being able to determine a cointegrating relationship between the rates. These contradicting results concerning the German and British spot exchange markets could be caused by the shorter data period used by Hakkio and Rush (1989), because cointegration is specifically a long-run relationship between the variables.

The third test of market efficiency deals with the unbiasedness hypothesis and examines within one country if the spot and forward rates are cointegrated. Hakkio and Rush report finding 'weak evidence' in favour of cointegration. As is clear from the work conducted by Baillie and Bollerslev (1989), merely establishing that the rates are

[17]The data consists of monthly spot and forward exchange rates for USD/GBP and USD/DEM for the period July 1975 until October 1986.

cointegrated does not prove that they are unbiased in their relation. For this purpose Hakkio and Rush formulate their fourth and fifth tests of market efficiency and estimate an error-correction equation:

(3.6)

$$(S_{t+1} - S_t) = c + a(S_t - dF_{t-1}) + b(F_t - F_{t-1}) + f(S_t - S_{t-1}) + g(F_{t-1} - F_{t-2}) + e_t$$

where S and F stand for spot and forward exchange rates, respectively (not in logarithmic form). In order to assert market efficiency should $-a = d = b = 1$ and $f = g = 0$. According to the results Hakkio and Rush are forced to reject their joint hypothesis of no predictable risk premium and efficient use of information by the market participants. They describe this as the strongest result of the study, but remain unconscious as to which factor (or factors) of the joint hypothesis causes the rejection.

The most recent paper on determining foreign exchange market efficiency based on the cointegration analysis is Clarida and Taylor (1993). The recent nature of the paper benefits from the development in the field of unit root econometrics. Clarida and Taylor develop a simple theoretical relationship between the spot and forward rates and show that the forward premia forecast the spot exchange rate. As the data they use weekly spot and 4, 13, 26, and 52 week forward exchange rates for the Deutschmark and the Pound Sterling against the US Dollar during period 1977:1 until 1990:26. They easily find the rates non-stationary. Then they examine efficiency by the simple efficiency hypothesis:

$$(3.7) \qquad s_{t+n} - s_t = \alpha + \beta(f_{n,t} - s_t) + \varepsilon_{t+n}$$

where $(s_{t+n}-s_t)$ is the change in the spot exchange rate at time (t+n) and $(f_{n,t}-s_t)$ is the corresponding forward premium. Like many researchers before them, Clarida and Taylor cannot confirm the null hypothesis that the forward premium is an unbiased predictor of the spot rate change, i.e. $\beta=1$. Because the rates are all integrated of order one, they are able to investigate their relations by an error-correction model. Clarida and Taylor use a dynamic vector error correction model (VECM) (Engle & Granger 1987 and Johansen 1991) for the system of the spot rate and the term structure of forward exchange rates:

$$(3.8) \qquad \Delta y_t = \mu + \Gamma_1 \Delta y_{t-1} + \cdots + \Gamma_{k-1} \Delta y_{t-k+1} + \Pi y_{t-k} + \varsigma_t$$

where $y_t = [s_t, f_{4,t}, f_{13,t}, f_{26,t}, f_{52,t}]'$ i.e. (j+1=) 5 by 1 vector of the rates and if the rank (r) of matrix Π is not full, but of reduced form, $r < 5$, the matrix can be factored into the product of two 5 by r matrices α and β:

$$(3.9) \qquad \Pi = \alpha\beta'$$

where β' is the r by 5 matrix of r cointegrating vectors and α is the 5 by r matrix of r adjustment coefficients for the five equations in the system.

Their framework, which draws upon Hall et al. (1992), Stock and Watson (1988), and Johansen (1991), predicts that in a system of (j+1) variables (here: j forward exchange rates and one corresponding spot rate) is exactly j (= [j+1]-[number of common trends]) cointegrating vectors and exactly one common trend, which is the non-stationary cointegrating parameter of each of the j forward rates and the one spot rate. The tests conducted support this. Further they test and accept the hypothesis that a basis for the space of the cointegrating relationships is the vector of the four (j = 4) forward premia: $[f_{4,t} - s_t, f_{13,t} - s_t, f_{26,t} - s_t, f_{52,t} - s_t]'$.This means that the cointegrating relationships can be expressed in the four-dimensional space in relation to the premia. From the above results Clarida and Taylor draw the conclusion that a VECM models the systems of spot exchange rate and the term structure of forward rates for both currencies very well. In order to test whether the term structure of the forward premia contains additional predictive content for the spot exchange rate, they examine the FIML[18] estimates of the five-equation VECMs for both the currencies. This enables them to conclude that, indeed, the spot rate is not exogenous with respect to lagged information in the term structure of the forward premia; the premia do contain statistically significant information about the future conduct of the spot exchange rate that is not contained in the lagged spot rate. This means that the spot rate is Granger-caused[19] by other information than its own history only. Clarida and Taylor still want to prove the usefulness of the information in the term structure of the forward rates and conduct out-of-sample forecasts, which show that the forecasting error of the spot rate can be reduced by using the information contained in the term structure of the forward premia by more than 33% at a six month horizon and by 50 to 90% at a one-year

[18]FIML is Full Information Maximum Likelihood for short.
[19]Granger causality between e.g. two variables means that one can help forecast the other (Engle & Granger 1991, 70).

horizon. It should be noted, however, that Clarida and Taylor do not maintain that the forward premium is an unbiased predictor of the spot exchange rate.

Our study follows mainly that of Clarida and Taylor (1993). There are some major differences, however. For instance, our analysis of unit roots in the time series under study is followed by model misspecification tests, which play an important role in model specification and are totally ignored by Clarida and Taylor. Furthermore, we record the numerical values of the cointegrating relationships, test the robustness of the Johansen cointegration tests, and conduct parameter restriction tests on the estimated cointegration space parameters. Similar aspects are not considered by Clarida and Taylor (op. cit.); they contend themselves with the result that the forward rates do contain relevant information in respect of future behaviour of the spot rate. We intend to find out more about the relationship.

4. ECONOMETRIC METHODOLOGY

4.1. Non-stationarity, cointegration, and error correction

A time series[20] (x_t) is said to be stationary if its mean, $E(x_t)$, is independent of t and its variance, $E[x_t-E(x_t)]^2$, is finite and does not vary with time. That is, they are both constant. A stationary series thus tends to return to its mean and fluctuate around it within a more-or-less constant range.

[20]A time series is a single occurrence of a random event (Greene 1993,413).

If a series must be differenced d times before it becomes stationary, it is said to be integrated of order d, denoted I(d). It follows, that if x_t is non-stationary and I(d), the d^{th} order differential of x_t, $\Delta^d x_t$, is stationary, and thus integrated of order zero, denoted $\Delta^d x_t \sim I(0)$.[21] (Hall et al. 1992, 130)

Consider two series y_t and w_t which are both integrated of order d, d > 0, and thus non-stationary. Their linear combination

(4.1) $z_t = y_t - \beta w_t$

is generally integrated of the same order d, too. But there are cases when the linear combination of two non-stationary time series is stationary. Here z_t is stationary: $z_t \sim I(d-b)$, when (d-b), where b>0, is zero and the mean of z_t is zero. Then it is said that y_t and w_t are cointegrated and the vector $(1, -\beta)$ is defined as the cointegrating vector. β being the cointegrating parameter. The space spanned by the cointegrating vectors is called the cointegration space. (Muscatelli & Hurn 1992, 2 and Hall et al. 1992, 117)

Co-integration can be understood as a long-run equilibrium relationship between the two series: the series move closely together and the difference between them is constant in the long run.

An important result of the co-integration analysis is the Granger Representation Theorem. (Granger 1986 and Engle & Granger 1987). According to the theorem there exists an error-correction representation of the data if the set of variables are

[21] $\Delta x_t = x_t - x_{t-1}$ and $\Delta^2 = \Delta(\Delta x_t)$ etc.

cointegrated of order d, d > 0. Actually, cointegration is a necessary condition for the existence of error correction representation and, equivalently, if the error correction representation is possible, it implies the variables are cointegrated (Engle & Granger 1991, 10).

Using the notations above, it is possible to write the error correction model (ECM) for bivariate case, i.e. a case with two variables, as follows:

$$(4.2) \qquad \begin{aligned} \Delta y_t &= A(L)\Delta y_{t-1} + B(L)\Delta w_{t-1} - \gamma_1 z_{t-1} + \varepsilon_{1t} \\ \Delta w_t &= C(L)\Delta y_{t-1} + D(L)\Delta w_{t-1} - \gamma_2 z_{t-1} + \varepsilon_{2t} \end{aligned}$$

where

$$z_t = y_t - \beta w_t$$

and A(L), B(L), C(L), D(L) are finite order lag polynomials with L as the lag operator[22], at least one of γ_1, γ_2 is non-zero and (ε_{1t}, ε_{2t}) is white noise[23] and may be correlated (Engle & Granger 1991, 10). All the variables in the above two ECMs are stationary. The ECM provides a way of separating the long run relationship between the economic variables ($y_t = \beta w_t$) from the short run responses (Δy_t, Δw_t terms). (Muscatelli & Hurn 1992, 2 and Engle & Granger 1991, 7) It is easy to see where the name error correction stems from: The components of the term z_{t-1} are deviations from the long run

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[22]For instance $A(L) = \alpha_0 - \alpha_1 L - \alpha_2 L^2 - \cdots - \alpha_p L^p$, is a p^{th} order autoregressive (AR) model, where for example $L^i y_t = y_{t-i}$, i = 1, . . ., p (Harris 1995, 3).

[23]A white noise process is a stationary process that has a zero mean, finite variance and is uncorrelated over time.

equilibrium. Therefore the system 'corrects the errors' from the equilibrium, when $0<\gamma_1<1$ and $0<\gamma_2<1$ (Pere 1990, 47).

4.2. The Engle-Granger two-step procedure in determining cointegrating vector

Engle and Granger (1987) introduced a two-step approach in determining the cointegrating vector for time series integrated of order one, I(1), and having a stationary, I(0), cointegrating relation. The first step of the procedure is to estimate the cointegrating regression in levels of the variables by regressing one of the variables on the other using OLS (Engle & Granger 1991, 9-11 and Banerjee et al. 1993, 157-161):

(4.3) $$y_t = \hat{\beta} w_t + z_t$$

where z_t are the residuals. The regression estimates on all the parameters are consistent and share a special feature characteristic of OLS estimates of cointegrated series: the estimates are superconsistent. Superconsistency means, that as more observations are added, i.e. $t = 1, \ldots, T$ and as $T \rightarrow \infty$, the estimates converge to their true values at the speed of T^{-1}, T being the number of observations, when in normal OLS regressions with stationary variables the rate of convergence is $T^{-\frac{1}{2}}$. It should be noted that it does not matter which way the above regression is run and neither is there a need for correction for simultaneous equations or serial correlation, because of the non-stationary, infinite variance, characteristic of the series.

The second step in the Engle-Granger procedure is to form an error correction model using the residuals, $z_t = y_t - \hat{\beta} w_t$, from (4.3) lagged:

$$\text{(4.4)} \qquad \begin{aligned} \Delta y_t &= A(L)\Delta y_{t-1} + B(L)\Delta w_{t-1} - \gamma_1 \left(y - \hat{\beta} w\right)_{t-1} \\ \Delta w_t &= C(L)\Delta y_{t-1} + D(L)\Delta w_{t-1} - \gamma_2 \left(y - \hat{\beta} w\right)_{t-1} \end{aligned}$$

All the terms in (4.4) are in changes, which are stationary, I(0), by assumption plus the error correction term, $y_t - \hat{\beta} w_t$. If y_t and w_t are cointegrated with $\hat{\beta}$ as the cointegrating parameter, then the error correction term is I(0), too and the orders of integration match on both sides of the equations. This can be verified by e.g. the Dickey-Fuller unit root tests, which will be introduced in section 5.1. Engle and Granger (1987) prove that if a superconsistent estimate of β such as $\hat{\beta}$ from the cointegratong regression is used, then the estimates for γ_i, i=1,2, and the coefficients for Δy_{t-1} and Δw_{t-1} will be as efficient asymptotically as if β were known in advance. (Engle & Granger 1991,10) The same cannot be said of the estimated standard errors of the model because the distribution of the OLS estimator is generally non-normal (Muscatelli & Hurn 1992, 3). Therefore the standard errors cannot for instance be used to test the significance of the regressors in the long-run, first stage equation or to assess if any restrictions should be imposed on the long run response, β. The choice of regressors to be included in the long-run equation has to be made before everything else. This brings in the potential of misspecification bias because the short run dynamics are excluded at the first stage. Engle and Yoo (1991) propose a third step to the procedure in order to make the distribution of the OLS estimator, β, normal. This may improve estimates of some of variables, but results are ambiguous. Another problem with the Engle-Granger two-step approach is the small sample bias: the small sample properties of the estimators may vary from their asymptotic properties. Furthermore, the Engle-Granger two-step procedure faces several shortcomings when the number of variables is more than two and there would be more than one cointegrating relationship. (Muscatelli & Hurn 1992, 4-5)

4.3. The Johansen ML method in determining cointegrating vector

To overcome problems associated with the Engle-Granger two-step procedure when there are more than two variables in a system, Johansen (1988a, 1988b, 1991) has developed a maximum likelihood (ML) method for determining the cointegrating vector(s). As compared to the Engle-Granger two-step procedure, the method is expected to give different results because of several reasons:

i) if $r > 1$, the Engle-Granger method will not generate consistent estimates of any of the significant cointegrating vectors.
ii) the Johansen method takes into account short-run dynamics on the process of ΔX_t, as well as any seasonal shifts in form of dummy variables.
iii) in the Johansen method the variables in X_t do not have to be arbitrarily normalised, because the method uses information from the equations of each of the X variables to obtain the estimates of $\hat{\beta}$, which do not depend on any normalisation.

The starting point for the Johansen method is a vector autoregressive (VAR) model with Gaussian errors. The model portrays a system $X_t = [x_{1t}, \ldots, x_{nt}]$, the variables of which are all non-stationary and $X_t \sim I(1)$.

(4.5) $$X_t = A_1X_{t-1} + \cdots + A_kX_{t-k} + \mu + \Psi D_t + \varepsilon_t, \qquad t = 1, \ldots, T$$

where $X_{k-1}, \ldots, X_0$ are given, A_i are nxn constant coefficient matrices, D_t are centered seasonal dummies to take account of any deterministic seasonality[24], and ε_t is independently, identically, and normally distributed n-dimensional vector, with zero mean and covariance matrix Ω: $\varepsilon_t \sim$ i.i.d. $N_n(0, \Omega)$. The number of lags, k, is chosen so as to ensure that the residuals are white noise, that the lagged variables $X_{t-k-1}, \ldots, X_{t-n}$ do not have any impact on Y_t. (Johansen & Juselius 1992, 215 and Muscatelli & Hurn 1992, 24)

In order to distinguish stationarity in the model induced by either linear combinations or by differencing, the vector error correction model (VECM) is useful:

(4.6) $$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \Psi D_t + \varepsilon_t, \; t = 1, \ldots, T$$

where $\Gamma_i = -(I - A_1 - \cdots - A_i), i = 1, \ldots, k-1$ and $\Pi = -(I - A_1 - \cdots - A_k)$, where I is the identity matrix. Like in the bivariate error correction model (4.2) it is easy to see from (4.6) the short and long run adjustment to changes in X_t: the estimates of $\hat{\Gamma}_i$ and $\hat{\Pi}$ respectively (Harris 1995, 77).

When looking at equation (4.6), the important thing to note is, that all other lagged values of X_t are in differences except for X_{t-k}. Because X_t is non-stationary[25], all the time differences of the system are stationary, i.e. integrated of order zero. Therefore, to

[24] Seasonal dummies are centered to make sure that they sum to zero over time and do not have an effect on the asymptotic distributions (Harris 1995, 81).
[25] X_t is assumed to be integrated of order one.

maintain equilibrium[26] between the right and the left hand side of the equation the term ΠX_{t-k} has to be stationary. Since X_{t-k} is integrated of order one and non-stationary, the parameter matrix Π must be of a form to make the term ΠX_{t-k} stationary. This can happen in three different cases:

i) all the variables in X_t are stationary
ii) Π is a matrix of zeros, implying no cointegration between the variables of X_t
iii) $\Pi = \alpha\beta'$ and there exists (n-1) cointegrating relationships $\beta' X_{t-k} \sim I(0)$, where α is the rate at which the ΔX_t returns to equilibrium[27] and β is the cointegrating vector.

There is only one case in the above, which is interesting and applicable to the current problem: number iii). In this case Π is a nxn matrix, which is the product of two nxr matrices α and β. In β there are $r \leq (n-1)$ cointegrating vectors, which means that r columns of β form r linearly independent stationary combinations of the variables in X_t and (n-r) columns form non-stationary vectors, common trends. Only the stationary cointegrating vectors can be taken into account for the equation (4.6) to be stationary. This means that the last (n-r) columns of α are effectively zero and can be ignored. The latter implies that when determining the number of cointegrating vectors, $r \leq (n-1)$, it is equivalent to testing for the number of columns which are zero in α. And because r is the number of independent stationary combinations of the variables in X_t and the matrix β, testing for cointegration is a test of determining the rank of Π, that is the number of linearly independent columns in Π. (Harris 1995, 79) The rank, r, can also be called the order of cointegration (Muscatelli & Hurn 1992, 25).

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[26] If the left hand side of an equation is I(0), hence, all the terms in the right hand side must also be I(0).
[27] ΠX_{t-k} is the error correction term in (3.6).

Reconsidering the three cases above in the context of the rank of Π:

i) Π has full rank, $r = n$

ii) Π has a rank of zero, $r = 0$

iii) Π has reduced rank, $r \leq (n - 1)$

Following the discussion above, it is evident then that the test for cointegration between the variables in X_t is a test of reduced rank in Π.

Johansen (1988a, 1988b) proposes a maximum likelihood approach using a reduced rank regression, which gives eigenvalues $\hat{\lambda}_1 > \hat{\lambda}_2 > \ldots > \hat{\lambda}_n$ and the corresponding eigenvectors $\hat{V} = (\hat{v}_1, \hat{v}_2, \ldots, \hat{v}_n)$. There are r elements in $\hat{V}$ which determine the linear combinations of stationary relationships and can be denoted $\hat{\beta} = (\hat{v}_1, \hat{v}_2, \ldots, \hat{v}_r)$. The separate $\hat{v}_i' X_t$[28], i = 1, 2, . . . , r, combinations are the cointegrating relationships, because they form high correlations with the stationary ΔX_t elements in (4.6). If they were non-stationary, they could not achieve such high correlations. This explains the interpretation of $\hat{\lambda}_i$: it measures how strongly the cointegrating relations $\hat{v}_i' X_t$ are correlated with the stationary part of the model. The rest (n-r) combinations $\hat{v}_i' X_t$, i = r+1, . . . , n, are non-stationary and uncorrelated with the stationary elements in (4.6) having eigenvectors $\hat{\lambda}_i$, i = r+1, . . . , n. (Harris 1995, 87)

To illustrate the above we derive the log-likelihood function for (4.6) from multivariate normal distribution (Banerjee et al. 1993, 262-265):

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[28] $X_t \sim I(1)$

(4.7)

$$L\left(\Gamma_1,\ldots,\Gamma_{k-1},\Pi,\Omega\middle|(X_1,\ldots,X_t)\right)=\frac{Tn}{2}\log(2\pi)-\frac{T}{2}\log|\Omega|-\frac{1}{2}\sum_{t=1}^{T}\varepsilon'_t\,\Omega^{-1}\varepsilon_t$$

where the lower case π is the ratio of the circumference of a circle to its diameter; $\pi \approx 3.14$. Then we first concentrate $L(\cdot)$ with respect to Ω and see that $\hat{\Omega} = T^{-1}\sum_{t=1}^{T}\varepsilon_t\varepsilon'_t$. To focus on the matrix Π, we concentrate $L(\cdot)$ with respect to $(\Gamma_1,\ldots,\Gamma_{k-1})$. Because $\{\Gamma_i\}$ are unrestricted, we correct for the short-run dynamics by taking out the effects of $(\Delta X_{t-1},\ldots,\Delta X_{t-k+1})$ and regressing ΔX_t and X_{t-k} separately on the right-hand side of (4.6) to obtain vectors of residuals R_{0t} and R_{kt} respectively:

(4.8)
$$\Delta X_t = \mu + \psi D_t + \sum_{i=1}^{k-1}\hat{\Gamma}\Delta X_{t-i} + R_{0t}$$
$$X_{t-k} = \mu + \psi D_t + \sum_{i=1}^{k-1}\tilde{\Gamma}\Delta X_{t-i} + R_{kt}$$

where, when letting $Q_t = (\Delta X'_{t-1},\ldots,\Delta X'_{t-k+1})'$,

$$\left(\hat{\Gamma}_1,\ldots,\hat{\Gamma}_{k-1}\right)=\left(\sum_{t=1}^{T}\Delta X_t Q'_t\right)\left(\sum_{t=1}^{T}Q_t Q'_t\right)^{-1}$$

(4.9) and

$$\left(\tilde{\Gamma}_1,\ldots,\tilde{\Gamma}_{k-1}\right)=\left(\sum_{t=1}^{T}\Delta X_{t-k} Q'_t\right)\left(\sum_{t=1}^{T}Q_t Q'_t\right)^{-1}$$

Now, the the concentrated likelihood function $L*(\Pi)$ depends only on the residuals $\{R_{0t}, R_{kt}\}$ and can be written as

$$(4.10) \qquad L*(\Pi) = K - \frac{T}{2}\log\left|\sum_{t=1}^{T}(R_{ot} - R_{kt})(R_{ot} - \Pi R_{kt})'\right|$$

Second moment matrices are then computed for the residuals $\{R_{0t}, R_{kt}\}$ and their cross-products, $S_{00}, S_{0k}, S_{k0}, S_{kk}$, where

$$(4.11) \qquad S_{ij} = T^{-1}\sum_{t=1}^{T} R_{it}R'_{jt} \qquad\qquad i, j = 0, k$$

As a result, from equation (4.10),

$$(4.12) \qquad L*(\Pi) = K - \frac{T}{2}\log\left|S_{00} - \Pi S_{k0} - S_{0k}\Pi' + \Pi S_{kk}\Pi'\right|$$

Because we earlier restricted Π to the form $\Pi = \alpha\beta'$, the equation (4.12) above can be written as:

$$(4.13) \qquad L*(\alpha,\beta) = K_0 - \frac{T}{2}\log\left|S_{00} - \alpha\beta' S_{k0} - S_{0k}\beta\alpha' + \alpha\beta' S_{kk}\beta\alpha'\right|$$

Then concentrating $L*(\alpha,\beta)$ with respect to α, gives an expression for the maximum likelihood estimate for α as a function of β and a further concentrated likelihood function depending only on β. When the maximum likelihood estimate of β is obtained, the estimates of all other unknown parameters as functions of the maximum likelihood estimate of β can be solved. Thus substituting $\hat{\alpha} = S_{0k}\beta(\beta' S_{kk}\beta)^{-1}$ into (4.13) yields:

$$(4.14) \qquad L^{**}(\beta) = K_1 - \frac{T}{2}\log\left|S_{00} - S_{0k}\beta(\beta' S_{kk}\beta)^{-1}\beta' S_{k0}\right|$$

To obtain the maximum likelihood estimate of β, the likelihood function $L^{**}(\beta)$ is maximized with respect to β. This is equivalent to:

$$(4.15) \qquad \text{Min}\left|S_{00} - S_{0k}\beta(\beta' S_{kk}\beta)^{-1}\beta' S_{k0}\right|$$

With the help of the variance-covariance matrix of R_{0t} and R_{kt}, the assumption that $|S_{00}| = 0$ and the normalization that $\beta' S_{kk}\beta = I$ the minimizing problem takes the form:

$$(4.16) \qquad \text{Min}\left|\beta'(S_{kk} - S_{k0}S_{00}^{-1}S_{0k})\beta\right| \qquad \text{subject to } \beta' S_{kk}\beta = I$$

which requires finding the saddle-point of the Lagrangian,

$$(4.17) \qquad \left|\beta'(S_{kk} - S_{k0}S_{00}^{-1}S_{0k})\beta\right| - \phi\left[\text{trace}(\beta' S_{kk}\beta - I)\right]$$

where ϕ is the Lagrangian associated with the constraint $\beta' S_{kk}\beta = I$. Now the minimization problem is a general eigenvalue problem, where the maximum likelihood estimate of β is obtained as the eigenvectors corresponding to the r largest eigenvalues from solving equation

(4.18) $$\left|\lambda S_{kk} - S_{k0}S_{00}^{-1}S_{0k}\right| = 0$$

which gives the n eigenvalues $\hat{\lambda}_1 > \hat{\lambda}_2 > \ldots > \hat{\lambda}_n$ and the corresponding eigenvectors $\hat{V} = (\hat{v}_1, \ldots, \hat{v}_n)$ which are normalised by $\hat{V}' S_{kk} \hat{V} = I$. (Banerjee et al. 1993, 261-265) Because the eigenvalues are the largest squared canonical correlations between the levels residuals R_{kt} and the difference residuals R_{0t}, the cointegrating vectors are

(4.19) $$\hat{\beta} = (\hat{v}_1, \hat{v}_2, \ldots, \hat{v}_r)$$

and the estimate of α follows from the definition $\Pi = \alpha\beta'$.

4.4 Testing for cointegration in the Johansen ML method

Johansen has developed two tests for assessing the number of cointegrating vectors in a system of variables.

The likelihood ratio test evaluates the hypothesis that matrix Π has reduced rank, $r \leq (n-1)$, versus the null of full rank, $\mathcal{H}_0$: $r = n$, which corresponds the VAR model in (4.5). The test statistic is as follows:

(4.20) $$-2\ln Q(\mathcal{H}_1(r)|\mathcal{H}_0) = -T\sum_{i=r+1}^{n} \ln(1-\hat{\lambda}_i) \qquad r = 0, 1, \ldots, n\text{-}1$$

where $H_1(r)$: $\Pi = \alpha\beta'$, r meaning the number of cointegrating vectors. The above test statistic is called the trace statistic and it is the multivariate analogue of the Dickey-Fuller test (see section 5.1) (Banerjee et al. 1993, 267). The other test statistic called the maximum eigenvalue test statistic (λ_{max}) is based on the comparison of $\mathcal{H}_1(r)$ versus $\mathcal{H}_2(r+1)$ which thus compares the alternative of r cointegrating vectors against (r+1) cointegrating vectors:

(4.21) $$-2\ln Q\left(\mathcal{H}_1(r)\middle|\mathcal{H}_2(r+1)\right) = -T\ln\left(1-\hat{\lambda}_{r+1}\right) \qquad r = 0, 1, \ldots, n\text{-}1$$

(Johansen 1991, 1555, Johansen & Juselius 1990, 175-176, and Johansen & Juselius 1992, 215-217) The critical values on both the test statistics are given for instance in Johansen (1988b), Johansen & Juselius (1990) and Osterwald-Lenum (1992).

It is essential to find the right rank; both under- and over estimation of the number of cointegration vectors, r, bring in potentially heavy impact for estimation and inference. On the one hand, under-estimation means that empirically relevant error correction terms have been omitted. The omitted terms are then passed on to the residual, e_t. On the other hand, over-estimation means that the distributions of statistics will be non-standard. This is because the term ΠX_{t-k} in (4.6) is not stationary, I(0), as it must be, if there is more than the correct number of cointegrating vectors. (Banerjee et al. 1993, 262)

5. THE DATA AND ANALYSIS

In finding out the relationship between the spot and forward exchange rates, we will use daily data on the FIM/USD and FIM/DEM spot rates[29] and one, three, and six month forward rates[30], which are calculated as an average of bid and ask quotes. The data starts six months after the date the markka was allowed to float, i.e. on 8 September, 1992, and covers thus the period of 9 March, 1993 until 29 April, 1994. The starting date was selected because the floating of the markka should provide an exchange rate that is the correct rate for the markka in the international foreign exchange market and thus affected only by the available information in the market.

The reason for conducting the research in two different exchange rates, FIM/USD and FIM/DEM, is simply to ascertain that the results hold in more than one currency. If the research was made in one exchange rate only, the results might not be as reliable, as they could be affected by the movement of the particular rate itself. It may be noted, however, that the crossrate between the US Dollar and the German Mark definitely affects the Finnish markka rate with the corresponding currencies.

The rates are daily observations and they amount to 291. For the reasons mentioned in section 2.3., logarithms have been taken of the rates. To avoid problems of matching the

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[29]The spot rate data is kindly provided by the Bank of Finland. Note, we use expressions FIM/USD and FIM/DEM, although it is more common to state the currency with more 'economic status' as the nominator. However, our expression is mathematically correct, when we consider the way the rates are quoted in our data.

[30]The forward rate data is kindly provided by Postipankki Oy.

spot and forward rates, only the starting dates were carefully calculated to be analogue[31].

5.1. Testing for unit roots

Before making any conclusions about cointegration between the spot and forward exchange rates, we must have evidence on the stationarity of the four time series. Therefore we are using the Augmented Dickey Fuller (ADF) unit root tests in The Shazam Econometrics Computer Program version 7.0 (Shazam 1993) and PcGive 8.0 (Doornik & Hendry 1994b). The reason why the tests are called augmented, is that they are modified from the 'original' Dickey Fuller tests[32] in a way that they take into account possible serial correlation in the time series. The factor which makes this possible is the term lagged values of the time series.

In the tests the time series are in logarithmic form. Two forms of the augmented Dickey-Fuller tests are conducted: a) with-constant ($\alpha_0 \neq 0$), no-trend and b) with-constant ($\alpha_0 \neq 0$), with-trend ($\alpha_2 \neq 0$). Adding a trend to the regression removes the possible effect of a trend to the time series. The idea thus is, that one should use the test regression the structure of which is closest to the data generating process of the time series under study. The two regression equations on a time series Y are constructed below:

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[31] Holidays and weekends cause a problem of matching spot and forward dates. The problem may either be forward contracts of different maturities maturing on the very same day or no forward contracts at all maturing on a particular day. From the point of view of this study, in the prior case there is an excess of estimations for the day's spot value and in the latter, no prediction at all. See e.g. Grabbe 1991,75-76.

[32] For instance of the sort: $\Delta Y_t = \alpha_0 + \alpha_1 t + (\alpha_2 - 1)Y_{t-1} + \varepsilon_t$ where $\varepsilon_t \sim IID(0, \sigma_\varepsilon^2)$, for $t=1, \ldots, T$ and in t-test H_0: $\alpha_2=1$.

$$\text{a)}\qquad \Delta Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta Y_{t-j} + \varepsilon_t$$

(5.1)

$$\text{b)}\qquad \Delta Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \alpha_2 t + \sum_{j=1}^{p} \gamma_j \Delta Y_{t-j} + \varepsilon_t$$

In the above regressions ε_t is white noise. This is ensured by adding enough lagged terms to the regression, which is done automatically by the econometric programme in use.

In the framework of the regressions the following tests are conducted on both the time series and the change in the series:

i) t-test, which is the test of a unit root in the time series, with H_0: $\alpha_1=0$
ii) F-test $\Phi_{1,2}$ which is a unit root test with zero drift in the time series, with a) H_0: $\alpha_0=\alpha_1=0$ and b) H_0: $\alpha_0=\alpha_1=\alpha_2=0$
iii) F-test Φ_3 which is a unit root test with non-zero drift, with b) H_0: $\alpha_1=\alpha_2=0$

The F-tests thus include multiple restrictions on the parameters. The critical values of all the tests are derived from the appropiate Dickey-Fuller distribution, not the standard t-distribution. In Shazam the asymptotic critical values are from Fuller (1976), Dickey and Fuller (1981), Guilkey and Schmidt (1989), and Davidson and MacKinnon (1993).

In PcGive the critical values originate from MacKinnon (1991)[33]. The results are reported in the tables below.

AUGMENTED DICKEY FULLER UNIT ROOT TEST RESULTS FOR FIM/DEM

	a) incl. constant (α_0) t-ratio H_0: $\alpha_1=0$	a) incl. constant (α_0) F-testΦ_1 H_0: $\alpha_0=\alpha_1=0$	b) incl. constant (α_0) and trend ($\alpha_2 t$) t-ratio H_0: $\alpha_1=0$	b) F-testΦ2 H_0: $\alpha_0=\alpha_1=\alpha_2=0$	b) F-testΦ3 H_0: $\alpha_1=\alpha_2=0$
s_t	-2.3164	3.0909	-2.6166	2.6880	3.6226
$f_{1,t}$	-2.3899	3.1292	-2.5750	2.6125	3.6447
$f_{3,t}$	-1.3424	0.90459	-1.8045	2.9727	4.4555
$f_{6,t}$	-1.7082	1.7163	-1.6175	1.1443	1.4600
Δs_t	-3.0301*	4.5919*	-3.0718	3.1870	4.7793
$\Delta f_{1,t}$	-3.8551*	7.4353*	-3.8483*	5.0378*	7.5522*
$\Delta f_{3,t}$	-2.8327*	4.0326*	-3.2644*	3.5907	5.3653*
$\Delta f_{6,t}$	-14.387*	103.52*	-14.355*	68.794*	103.16*
critical values at 10% significance level	-2.57	3.78	-3.13	4.03	5.34

* indicates rejection of the H_0 at 10% significance level.

Table 5.1

From table 5.1 it is easy to see that at 10% significance level the null hypothesis of non-stationarity cannot be rejected for the FIM / DEM time series in levels. As to the first differences of the time series, the null is rejected in every case. One interesting feature there is, however, in the tests for the first difference of the spot exchange rate. When testing for a unit root in the context of the Augmented Dickey Fuller unit root test with

[33]MacKinnon uses the technique of response surface regressions to estimate critical values for unit root tests. This method enables him e.g. to estimate asymptotic critical values without using infinitely large samples. (MacKinnon, 1991)

constant and trend, the results are in favour of the null hypothesis of a unit root. The test statistics are very close to the critical values, though. In order to acquire comparative results, we conducted the Augmented Dickey Fuller test on PcGive econometric programme.

AUGMENTED DICKEY FULLER UNIT ROOT TEST RESULTS FOR FIM/DEM

	a) incl. constant (α_0) t-ADF H_0: α_1=0	lag	**b)** incl. constant (α_0) and trend ($\alpha_2 t$) t-ADF H_0: α_1=0	lag
s_t	**-1.6511**	**2**	**-1.8445**	**2**
	-1.6898	1	-1.9116	1
	-1.5462	0	-1.6314	0
$f_{1,t}$	**-1.2179**	**2**	**-2.0406**	**2**
	-1.2971	1	-2.1331	1
	-1.4404	0	-2.2905	0
$f_{3,t}$	-2.9976**	2	-3.2707	2
	-4.0598***	1	-4.3107***	1
	-6.6721* **	**0**	**-6.9230*** **	**0**
$f_{6,t}$	**-1.7527**	**2**	**-1.7590**	**2**
	-1.7673	1	-1.7822	1
	-1.8816	0	-1.9456	0
critical value at 5% significance level	-2.872		-3.427	
critical value at 1% significance level	-3.455		-3.993	
Δs_t	-9.2349***	2	-9.2355***	2
	-112.256***	1	-11.252***	1
	-14.054* **	**0**	**-14.043*** **	**0**
$\Delta f_{1,t}$	-11.445***	2	-11.413***	2
	-13.528***	1	-13.494***	1
	-19.521* **	**0**	**-19.484*** **	**0**
$\Delta f_{3,t}$	-15.959***	2	-16.032***	2
	-19.730***	1	-19.761***	1
	-28.308* **	**0**	**-28.298*** **	**0**
$\Delta f_{6,t}$	-14.819***	2	-14.795***	2
	-14.387***	1	-14.355***	1
	-18.566* **	**0**	**-18.535*** **	**0**
critical value at 5% significance level	-2.872		-2.872	
critical value at 1% significance level	-3.455		-3.993	

** indicates rejection of the H_0 at 5% significance level.
*** indicates rejection of the H_0 at 1% significance level.

Table 5.2

In PcGive the ADF unit root test are conducted in a slightly different way. We only examine the ADF t-test including a) a constant and b) a constant and trend in the regressions. PcGive studies the unit root by calculating t-values for lags between zero and two, which is equivalent to p=0, 1, 2 in equation (5.1).[34] The results are presented in table 5.2 above, with the appropriate lag row highlighted for each variable.

The first difference of the spot exchange rate for FIM/DEM is easily found stationary in table 5.2. All the other results follow the pattern of table 5.1 except for the three month forward exchange rate, which is found to be stationary at one percent significance level. This is not impossible, and, in fact, when looking at the graph of the three month forward exchange rate for FIM / DEM over the study period (see Appendix 2), it seems visually quite stationary. The fact that one time series in a vector, where all the other series under investigation are non-stationary, is stationary, does not restrain our analysis of cointegration in the next section of the study. According to Engle and Granger (1991, 14) and in conjunction with Johansen (1991, 1552) including a stationary variable in cointegrating relationship should not affect the remaining variables, nor should it have an effect on the asymptotic critical values of the test statistics. The interpretation and implications are discussed in more detail in section 5.4.1. If, however, one series in the midst of other I(1) series was integrated of order two, then it would not be so straightforward to proceed with the testing.

As to the FIM / USD exchange rates, the results of the Augmented Dickey Fuller unit root tests are shown in tables 5.3 and 5.4. for levels and first differences. The time series seem to be non-stationary without differencing and stationary after first

[34]Naturally, given the nature of the ADF tests, lags are calculated in Shazam, too. But in Shazam the programme automatically chooses the number of lags needed.

differences. One exception is found again, however. The time series for the three month forward exchange rate is stationary in the test regression a) with constant, without trend. The result is similar to the one in table 5.1 in regard to the first difference of the FIM / DEM spot exchange rate. In both cases the null is rejected when a constant is added to the ADF regression without the trend. This suggests that in these two cases the particular time series is generated by a process which is similar to equation (5.1) a). According to Harris (1995, 31-32) a sequential testing procedure with respect to the two different forms of Augmented Dickey Fuller test is possible. This means that testing is started with the general form regression which includes all possible nuisance parameters, like the constant and trend, and is then carried on by eliminating the nuisance parameters step by step until the null hypothesis is rejected. Here it means starting from the regression presented by equation (5.1) b) and then moving on to (5.1) a), which implies then, that the series are stationary.

AUGMENTED DICKEY FULLER UNIT ROOT TEST RESULTS FOR FIM/USD

	a) incl. constant (α_0)		b) incl. constant (α_0) and trend ($\alpha_2 t$)		
	t-ratio H_0: $\alpha_1=0$	F-testΦ_1 H_0: $\alpha_0=\alpha_1=0$	t-ratio H_0: $\alpha_1=0$	F-testΦ2 H_0: $\alpha_0=\alpha_1=\alpha_2=0$	F-testΦ3 H_0: $\alpha_1=\alpha_2=0$
s_t	-2.3932	3.1887	-2.4187	2.1660	2.9251
$f_{1,t}$	-2.4424	3.1291	-2.4964	2.1757	3.1174
$f_{3,t}$	-2.8087*	4.0642*	-2.4580	2.7826	4.0544
$f_{6,t}$	-2.0294	2.5362	-1.9657	1.8812	2.3457
Δs_t	-3.7571*	7.0593*	-3.7417*	4.6887*	7.0316*
$\Delta f_{1,t}$	-3.7046*	6.8645*	-3.6955*	4.5588*	6.8356*
$\Delta f_{3,t}$	-6.1555*	18.948*	-6.3929*	13.642*	20.460*
$\Delta f_{6,t}$	-5.7185*	16.385*	-5.7550*	11.204*	16.772*
critical values at 10% significance level	-2.57	3.78	-3.13	4.03	5.34

* indicates rejection of the H_0 at 10% significance level.

Table 5.3

In table 5.4 the results given by PcGive indicate yet another time series stationary without differencing: the one-month forward exchange rate for FIM / USD. The reason for the observation that in t-tests conducted in PcGive, there are more stationary time series in levels than in Shazam, draws upon the differences in the testing techniques and the used significance levels in the programmes. In PcGive the significance levels reported are five and one percent, in Shazam ten percent. We found it interesting to see the effect of different significance levels on the data.

The results concerning the unit root tests indicate that the spot and forward exchange rates are non-stationary with possibly few exceptions. The first differences are clearly all stationary. These conclusions coincide with earlier results of similar studies. For instance Coleman (1990) found among other currencies (see footnote 16) the logarithm of the spot exchange rate of the Finnish Markka against the US Dollar non-stationary. MacDonald and Taylor (1989) and Clarida and Taylor (1993) are able to draw the same conclusion on several currencies (see pages 17, 19, respectively).

An interesting perception can be made from the unit root tests which record the number of lags used. The first differences of the time series in both the US Dollar and the Deutschmark are stationary at one percent level without any lags in the ADF test. This means that they are best described by a simple random walk process, which does not seem to be true for the time series in levels.

AUGMENTED DICKEY FULLER UNIT ROOT TEST RESULTS FOR FIM/USD

	a) incl. constant (α_0) t-ADF H_0: α_1=0	lag	**b)** incl. constant (α_0) and trend ($\alpha_2 t$) t-ADF H_0: α_1=0	lag
s_t	**-2.3574**	**2**	**-2.3655**	**2**
	-2.3959	1	-2.4054	1
	-2.4059	0	-2.4152	0
$f_{1,t}$	-2.6559	2	-2.8152	2
	-3.4690***	1	-3.6380**	1
	-5.9086* **	**0**	**-6.1443*** **	**0**
$f_{3,t}$	-3.1238**	2	**-2.6677**	**2**
	-3.1936**	1	-2.7910	1
	-3.4658* **	**0**	-3.2080	0
$f_{6,t}$	**-2.3908**	**2**	**-1.9924**	**2**
	-2.4029	1	-1.9959	1
	-2.4183	0	-2.1602	0
Δs_t	-10.593***	2	-10.578***	2
	-12.295***	1	-12.277***	1
	-17.065* **	**0**	**-17.038*** **	**0**
$\Delta f_{1,t}$	-14.708***	2	-14.684***	2
	-19.032***	1	-19.000***	1
	-28.999* **	**0**	**-28.948*** **	**0**
$\Delta f_{3,t}$	-10.892***	2	-11.077***	2
	-14.344***	1	-14.514***	1
	-21.484* **	**0**	**-21.617*** **	**0**
$\Delta f_{6,t}$	-13.047***	2	-13.183***	2
	-13.916***	1	-13.997***	1
	-18.820* **	**0**	**-18.904*** **	**0**
critical value at 5% significance level	-2.872		-3.427	
critical value at 1% significance level	-3.455		-3.993	

** indicates rejection of the H_0 at 5% significance level.
*** indicates rejection of the H_0 at 1% significance level.

Table 5.4

5.2. Power of unit root tests

The power of a test means its ability to reject a false null hypothesis. For the unit root tests discussed in the previous section the power is commonly recognised to be quite low. The reasons include:

i) the value of the autoregressive parameter α_1 in equation (5.1) under alternative hypothesis, which critically affect the power of the unit root tests when it is less than but close to unity (Muscatelli & Hurn 1992, 8-9).
ii) the small sample bias.
iii) the number of deterministic regressors in the test regression.

These are discussed in more detail below.

The small sample bias is pronounced in finite samples and in small numbers of observations. The unit root tests are then biased towards accepting the null hypothesis of non-stationarity when the data is close to having a unit root but is in fact stationary. (Harris 1995, 27) In addition to the size of the data, the sampling frequency has an effect on the power of the test. When the data is sampled more frequently over the period of interest, the power of the test increases only slightly. In fact it can be shown that data sets containing less observations over a longer time period have a higher test power than data sets which contain more frequent observations over a shorter period. The problem with very long observation periods is, that there might be a structural change included in the period. A structural change biases the test in favour of the unit root hypothesis, when, in reality, it should be rejected. (Campbell & Perron 1991, 153)

Above, we discussed the two different Augmented Dickey Fuller test regressions and which one of them should be used under different circumstances. It was noted that there should be as many nuisance parameters, constant and trend, in the regression as there are those in the data generating process of the time series under study or the power of the test will lose finite sample power. On the other hand, too many deterministic regressors in the test regression decrease the power of the unit root test. Another point about the ADF test regressions is the choice of the lag parameter p in equation (5.1). Too few lags have an effect on the size of the tests, resulting in over-rejecting the null hypothesis of a unit root when it is true. (Harris 1995, 34 and Campbell & Perron 1991, 151, 154)

Since there seems to be so many instances when the power of the ADF tests is decreased, an alternative test has been developed by Phillips and Perron (1988). Instead of adding more parameters, like a constant and trend, to the test regression, Phillips and Perron undertake a non-parametric correction to the t-test statistic for serial correlation and potential heteroskedosticity. This gives the test a substantially higher size-adjusted power than in the ADF tests. However, the Phillips-Perron tests can be less reliable on the whole. (Campbell & Perron 1991, 156, Harris 1995, 33, and Clarida & Taylor 1993, 12)

One of the ground reasons for the unrealiability of the unit root tests lies in the fact, that trend- and difference stationary series are extremely difficult to distinguish. 'In finite samples, any trend-stationary process can be approximated arbitrarily well by a unit root process (in the sense that the autocovariance structures will be arbitrarily close)'

(Campbell & Perron 1991, 157) and vice versa. This means that a high power unit root test against any stationary alternative will inevitably have equivalently high probability to reject the null hypothesis of a unit root falsely when used in near-stationary series.

Given the discussion above, our choice to use the Johansen-framework in the study of cointegration instead of the Engle-Granger procedure is even more understandable: the Engle-Granger procedure relies heavily on the Dickey Fuller unit root tests.

5.3. Residual misspecification tests

The Johansen procedure for determining the cointegrating vectors relies on a vector autoregressive (VAR) model with Gaussian errors as was introduced in section 3.3. Therefore we must make sure our data sets comply with the assumption. For both the currencies the system $X_t = [s_t, f_{1,t}, f_{3,t}, f_{6,t}]'$ is a four-dimensional vector autoregressive model:

$$X_t = A_1 X_{t-1} + \cdots + A_k X_{t-k} + \mu + \Psi D_t + \varepsilon_t, \qquad t = 1, \ldots, T \tag{3.5}$$

where $X_{k-1}, \ldots, X_0$ are given, A_i are nxn constant coefficient matrices, D_t are centered seasonal dummies, and ε_t~i.i.d. $N_n(0, \Omega)$.

The misspecification tests are run on the vector error correction form of the systems:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \Psi D_t + \varepsilon_t, \; t = 1, \ldots, T \tag{3.6}$$

where $\Gamma_i = -(I - A_1 - \cdots - A_i), i = 1, \ldots, k-1$ and $\Pi = -(I - A_1 - \cdots - A_k)$, where I is the identity matrix.

In the tests the maximum lag length, k, is chosen so as to ensure that the residuals are uncorrelated. This particular maximum lag length is then used in the cointegration tests.

The test results for vector error autocorrelation and vector normality for both the FIM/DEM and FIM/USD time series are listed below. The tests are conducted in PcFiml as a Lagrange multiplier form of the χ^2 test for vector error autocorrelation, and transformed residual test of skewness and kurtosis for vector normality. The F-form critical values in the autocorrelation test corresponds the small sample correction of the test statistic. The values in parentheses following the test statistics are the probability values of acceptance of the null hypothesis.

Model misspecification tests for FIM/DEM

k	Autocorrelation tests $\chi^2(32)$ H_0: no autocorrelation	F-Form(32,r)	Normality test $\chi^2(8)$ H_0: normality
3	48.308(0.0322)**	r = 949 1.4376(0.0560)	4485.8(0.0000)***
4	**36.029(0.2855)**	**r = 934** **1.0407(0.4064)**	**6066.6(0.0000)*****
5	30.129(0.5615)	r = 919 0.85074(0.7052)	7131.4(0.0000)***

** indicates rejection of the H_0 at 5% significance level.
*** indicates rejection of the H_0 at 1% significance level.

Table 5.5

In accordance with table 5.5 above, the conclusion can be drawn that the maximum number of lags to be included in the FIM/DEM vector error correction model (equation 3.6) and in the subsequent cointegration tests is a maximum lag of four periods. It is the first number of maximum lags which makes the residual autocorrelation disappear. In fact, lags five, six, and seven are also free of autocorrelation (k = 6, 7 not recorded in table 5.5), and we are conducting the cointegration tests also on five and six maximum lags (see section 5.4.3.). As to the vector normality tests, we cannot accept vector normality in any case. It is maintained by many researches (e.g. Boothe & Glassman 1987), that a time series for an exchange rate is rarely normally distributed. The distribution is often characterised by kurtosis and/or leptokurtosis. This can be quite well seen in Figure 5.1 in Appendix 4.

Model misspecification tests for FIM/USD

	Autocorrelation tests		Normality test
	$\chi^2(32)$	F-Form(32, r)	$\chi^2(8)$
k	H_0: no autocorrelation		H_0: normality
		r = 967	
2	49.45(0.0252)**	1.5142(0.0343)**	1088(0.0000)***
		r = 953	
3	**40.06(0.1550)**	**1.1962(0.2110)**	**2284.6(0.0000)*****
		r = 938	
4	59.793(0.0021)***	1.7925(0.0047)***	2786.2(0.0000)***

** indicates rejection of the H_0 at 5% significance level.
*** indicates rejection of the H_0 at 1% significance level.

Table 5.6

In the FIM/USD case the first lag length with no vector error autocorrelation is k = 3. The one step lower and higher lags bring both about autocorrelation in the residuals, unlike in the FIM/DEM case. Figure 5.2 in Appendix 4 shows residual correlograms for the FIM/USD VECM with k = 3. No residual serial correlation seems to exist.

Normality is again a prerequisite hard to fill. However, it can be noted that Johansen & Juselius (1990, 176) are willing to accept non-normally distributed data into their cointegration analysis with a remark that probably accepting data with excess kurtosis is less serious than accepting data with a skewed distribution. Some other studies, too, which are discussed in section 5.4.2., have come to the same conclusion.

5.4. Cointegration tests

In our analysis we are concentrating on a model with a constant and centered seasonal dummies, like the VECM in equation (3.6). Earlier in the theory of cointegration a model without those terms was preferred (Johansen 1988b), but further study has revealed that the constant is vital in interpreting the model and for the statistical and probabilistic analysis (Johansen 1991, 1551). As there are no linear trends in our series, we use a constant that is restricted to the cointegration space, i.e. $\mu = \alpha\beta_0'$. This does not alter the Johansen procedure for determining the cointegration rank as introduced in section 4.3 significantly and the results derived hold for the restricted constant alike.

Because the data is entered on a daily basis seasonal dummies are included in the analysis to take account of any seasonal shifts in the data. It has been maintained in many studies that a day-of-the-week effect is apparent in exchange rate time series (e.g. Baillie & Bollerslev, 1989 and Copeland & Wang, 1993) For example, the dollar is commonly known to follow a weekly cycle, where it weakens between Wednesday and Thursday close rates (Copeland & Wang 1993, 4). In the analysis the dummies are centered so that their mean is zero. In this way they do not have an effect on the fundamental distribution of the test statistics and the critical values given by the PcFiml econometric programme are valid. In fact, the cointegration tests were first conducted without any seasonals and the results were alarming already in the model misspecification tests. Thus a model with seasonals seems to fit the data better.

The results of the two conducted tests based on the Johansen procedure on determining the cointegrating vectors are reported in tables 5.7 and 5.8. The first column of the tables shows the rank, r, of the vector Π. The second column exhibits the estimated eigenvalues, $\hat{\lambda}_i$. The maximal eigenvalue test is reported then in the first set of rows. The alternative hypothesis against the null of $\mathcal{H}_1$: rank = r, is $\mathcal{H}_2$: r = r+1. The test statistic as it was introduced in section 4.4:

$$(4.21) \qquad -2\ln Q(\mathcal{H}_1(r)|\mathcal{H}_2(r+1)) = -T\ln(1-\hat{\lambda}_{r+1}) \qquad r = 0, 1, \ldots, n-1$$

The maximal eigenvalue test statistic is followed by two different critical values. The first is based on the equation (4.21) presenting the 95% confidence level, the second represents a small-sample correction, when the T in equation (4.21) has been replaced by (T-nk).[35] The small-sample correction is advocated by Reimers (1992) who found that the Johansen procedure rejects the null hypothesis too often in small samples. It is still uncertain whether the correction is for the better.

On the second set of rows of both the tables the same rank and eigenvalue information is presented as on the first set. This is followed by the trace statistic with similar critical values as for the maximal eigenvalue test. The trace statistic is repeated below from section 4.4:

[35] $t = 1, 2, \ldots, T$, n is the number of variables in the system, and k is the maximum number of lags in the VAR.

$$(4.20) \qquad -2\ln Q(\mathcal{H}_1(r)|\mathcal{H}_0) = -T\sum_{i=r+1}^{n}\ln(1-\hat{\lambda}_i) \qquad r = 0, 1, \ldots, n\text{-}1$$

The null hypothesis of r = n, meaning that the rank is equals the number of variables in the system or that matrix Π has full rank, is tested against $r \leq (n-1)$, that matrix Π has reduced rank.

TESTS OF THE COINTEGRATION RANK OF THE FIM/DEM DATA, k=4

H_0: rank = r	$\hat{\lambda}_i$	$-T\ln(1-\hat{\lambda}_{r+1})$	$\hat{\lambda}_{max}(0.95)$	$\hat{\lambda}_{max}(T\text{-}nk)$
r = 0		28.69**	28.1	27.08
$r \leq 1$	0.0957604	18.74	22.0	17.69
$r \leq 2$	0.0636471	13.89	15.7	13.11
$r \leq 3$	0.0475595	3.095	9.2	2.921

H_0: rank = r	$\hat{\lambda}_i$	$-T\sum\ln(1-\hat{\lambda}_i)$	$\hat{\lambda}_{trace}(0.95)$	$\hat{\lambda}_{trace}(T\text{-}nk)$
r = 0		64.41***	53.1	60.8***
$r \leq 1$	0.0957604	35.73**	34.9	33.72
$r \leq 2$	0.0636471	16.98	20.0	16.03
$r \leq 3$	0.0475595	3.095	9.2	2.921

** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 5% significance level.

*** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 1% significance level.

Table 5.7

The results in table 5.7 above for the FIM/DEM time series indicate that according to both the tests the null of no cointegrating vectors can be rejected in favour of one cointegrating vector. However, the trace statistic rejects even the rank one suggesting to accept rank two, but with a test statistic that is very close to the critical values. This illustrates the fact that the two tests, maximal eigenvalue and trace statistic, can give different results. The reason is basically the low power when the cointegrating relation is proximate to the non-stationary boundary (Johansen & Juselius 1992, 221). Graphical illustrations of the cointegrating relationships are useful in situations when there is ambiguity about the number of cointegrating vectors. Figure 5.3 in Appendix 5 illustrates the estimated relations, $\hat{\beta}_i'X_t$, i = 1, 2, 3, 4. In the figure, the cointegrating relations are shown as stationary. When taking the figure into account, the right answer to the question of the rank of the matrix Π seems to be one. The second vector is very close to being stationary, though.

TESTS OF THE COINTEGRATION RANK OF THE FIM/USD DATA, k=3

H_0: rank = r	$\hat{\lambda}_i$	$-T\ln(1-\hat{\lambda}_{r+1})$	$\hat{\lambda}_{max}(0.95)$	$\hat{\lambda}_{max}(T-nk)$
r = 0		28.89**	28.1	27.68
$r \leq 1$	0.0957462	21.04	22.0	20.16
$r \leq 2$	0.0706981	5.271	15.7	5.05
$r \leq 3$	0.0181969	3.33	9.2	3.191

H_0: rank = r	$\hat{\lambda}_i$	$-T\sum \ln(1-\hat{\lambda}_i)$	$\hat{\lambda}_{trace}(0.95)$	$\hat{\lambda}_{trace}(T-nk)$
r = 0		58.53**	53.1	56.08**
$r \leq 1$	0.0957462	29.64	34.9	28.4
$r \leq 2$	0.0706981	8.601	20.0	8.241
$r \leq 3$	0.0181969	3.33	9.2	3.191

** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 5% significance level.

*** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 1% significance level.

Table 5.8

The results concerning the FIM/USD data as shown in table 5.8 lead us to accept one cointegrating vector. The maximal eigenvalue statistic is in ranks zero and one very close to the test statistics. Should we accept two cointegrating vectors or, on the other hand, none at all? Figure 5.4 in Appendix 5 sheds light on the problem: it seems that we should accept one cointegrating vector or none at all. The first cointegrating relationship, $\hat{\beta}_1'X_t$, could be judged stationary, but not the rest. Therefore, on the basis of formal testing and graphical analysis our conclusion for the FIM/DEM data is one cointegrating vector in the system of the four variables and for the FIM/USD data one or none cointegrating vectors. Next we will look at the coefficients of the obtained vectors and find out if they can offer a solution to the FIM/USD data problem.

5.4.1 Cointegrating relations

The estimated coefficients of the cointegrating relations, $\hat{\beta}_i$, and the adjustment parameters, $\hat{\alpha}_i$, are shown in table 5.9 below.

	Standardised β' eigenvectors			
	FIM/DEM		FIM/USD	
	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_1$	$\hat{\beta}_2$
s_t	1.000	-1.401	1.000	-0.5951
$f_{1,t}$	-1.190	1.000	-1.624	1.000
$f_{3,t}$	-0.09035	0.1736	0.1134	1.214
$f_{6,t}$	-0.02195	1.382	0.05387	0.3248
μ	0.3733	-1.431	0.7946	-3.406
	Standardised α coefficients			
	$\hat{\alpha}_1$	$\hat{\alpha}_2$	$\hat{\alpha}_1$	$\hat{\alpha}_2$
s_t	0.007398	0.01129	0.006924	0.01908
$f_{1,t}$	0.06195	0.004660	0.08681	-0.04910
$f_{3,t}$	-0.03887	-0.03278	-0.04362	-0.01894
$f_{6,t}$	0.02305	-0.03381	-0.009755	-0.02513

Table 5.9

Since we decided on the basis of formal testing and graphical analysis that the correct number of cointegrating vectors for the FIM/DEM data is one, we are looking at the first columns of the estimated eigenvectors and adjustment parameters in table 5.9. The error correction, long-run, equilibrium relation between the variables of $X_t = [s_t, f_{1,t}, f_{3,t}, f_{6,t}]'$ is thus

(5.2) $$s_t = 1.19f_{1,t} + 0.09f_{3,t} + 0.02f_{6,t} - 0.37$$

and the corresponding adjustment parameters are

(5.3) $$\hat{\alpha}' = (0.007, 0.062, -0.389, 0.023)$$

Equation (5.2) is the statement for the long run equilibrium relationship between the spot exchange rate and the corresponding forward exchange rates. We must note, however, that the estimated coefficients are only indicative. It seems that the one-month forward rate has the strongest impact on the spot rate, the three- and six- month rates only adjust its effect on the spot rate. In fact the graph in Appendix 2 supports this finding: of all the forward rates it is the one-month rate that moves very much in accordance with the spot rate. The estimated coefficients are all of the 'correct' sign in respect of the interest rate parity and its implication to the efficient markets hypothesis, so that the spot and forward rates move in the dame direction. For instance if the forward rates all increase by 10 percent, the spot rate increases by 13 percent minus the value of the constant which results in an overall decrease of 24 percent. It seems that the estimates of the forward rates 'overestimate' the value of the spot rate and the constant then brings this estimate down to its correct level. Actually the second cointegrating vector, $\hat{\beta}_2$, gives results roughly similar to the first: the signs of the estimates are all 'correct' but the estimate of the six-month forward rate is almost unity and it would then have stronger impact on the spot rate than the other rates.[36] This is merely a curiosity, since we have already maintained the existence of a unique cointegrating vector.

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[36] $s_t = 0.71f_{1,t} + 0.12f_{3,t} + 0.99f_{6,t} - 1.02$

The adjustment coefficients of $\hat{\alpha}_1$ are all very close to zero for the FIM/DEM data. They can be given an economic meaning of the average speed of adjustment towards the estimated equilibrium state: a low coefficient indicates slow adjustment and a high coefficient instant adjustment (Johansen & Juselius 1990, 183) All the coefficients, as given in equation (5.3), are very small. Furthermore, the sign of the coefficient corresponding to the three-month forward rate is negative, while all the others are positive. It will be formally tested in the next section whether the coefficients one at a time are effectively zero.

Table 5.9 also exhibits the corresponding results for the FIM/USD analysis. It seems that the estimated long-run equilibrium relationship and the estimated adjustment parameters are the following:

$$(5.4) \qquad s_t = 1.62 f_{1,t} - 0.11 f_{3,t} - 0.05 f_{6,t} - 0.79$$

$$(5.5) \qquad \hat{\alpha}' = (0.007, 0.087, -0.044, -0.010)$$

The numerical values are not very different from those of the FIM/DEM data estimates, and the one-month forward rate seems to follow the spot rate closer than the other forward rates, which is evident in the graph of Appendix 3, but there is a striking difference: the signs of the forward rates are both positive and negative. This does not totally contradict the efficient markets hypothesis, because its interpretation with several forward rates as predictors of the spot rate is not clear. But, as stated before, the aim of this study is not to assert the existence of efficient markets as it is to maintain that the forward rates include information that is essential in the process of determining the future spot rate. Also the adjustment parameters are very small.

The role of the eigenvalues deserves some discussion. Because of the eigenvalues all the cointegration tests conducted above are highly dependent on one another. This follows from the fact that the eigenvalues are ordered, $\hat{\lambda}_1 > \hat{\lambda}_2 > \ldots > \hat{\lambda}_n$, and thus not independent. (Johansen & Juselius 1990, 181) The ordering of the eigenvalues also has an effect on the cointegration relations, so, that the first relation, $\hat{\beta}_1 X_t$, is the most correlated with the stationary part of the process ΔX_t and the second is the next most correlated, etc. There is also interdependence between the estimated coefficients of the cointegrating analysis and the eigenvalues as can be derived from the analysis in section 4.3 . (Johansen & Juselius 1992, 221) In our case the low values of the eigenvalues in both the FIM/DEM and FIM/USD data are definitely affected by the values of $\hat{\alpha}_i$ and $\hat{\beta}_i$, most of which are close to zero.

When we discussed the results of the unit root tests in section 5.1, we noticed that according to the results it seemed that in both the systems there were possibly some stationary time series even without differencing. These were the three-month forward rate for the FIM/DEM series and the one- and three-month series for the FIM/USD. Now, according to the cointegration tests conducted above, this seems not to be the case. A stationary variable among non-stationary variables would show in the cointegration analysis as an additional cointegrating vector and some cointegrating vectors in the matrix β of cointegration vectors, should contain only one variable, because the stationary variable forms a linearly independent column in matrix $\Pi(= \alpha \beta')$ by itself. (Harris 1995, 80)

Given the results of the cointegration tests, we wanted to make some adjustments to the original model. What if the model is better described by only two variables, say the spot rate and the one-month forward rate? Or the spot rate, one-month forward rate, and the six-month forward rate? We conducted the cointegration tests on seven different possibilities:

$$X_{1t} = [s_t, f_{1,t}]' \qquad X_{4t} = [s_t, f_{1,t}, f_{3,t}]'$$

$$X_{2t} = [s_t, f_{3,t}]' \qquad X_{5t} = [s_t, f_{1,t}, f_{6,t}]'$$

$$X_{3t} = [s_t, f_{6,t}]' \qquad X_{6t} = [s_t, f_{3,t}, f_{6,t}]'$$

with very slim results. The reason is clear: when the variables in the system change, the whole model changes and for instance the original model misspecification tests do not apply. The solution is to test different combinations of variables in the long run error correction equation and to do it by parameter restriction tests, unless, of course, you want to change your model specification altogether. We find it best to continue the analysis with the original model and to test cointegration test induced hypotheses about the variables by formal parameter restriction tests in section 5.4.4.

5.4.2. Power and finite sample properties of the Johansen ML cointegration tests

Not much is known about the power properties of the Johansen tests. Johansen (1989, as quoted in Banerjee et al. 1993, 277) has studied the power function of the trace statistic from the point of view of 'near integrated' processes[37]. He found that the ability of the test to reject a false null hypothesis falls when (n - r) rises, because a larger cointegrating space has to be searched in order to find the cointegrating vectors. The power also depends on the magnitude of the ECM impact and on the position of the 'local' cointegrating vectors in the space. Given the discussion above, care should be taken when interpreting the cointegration test results. Therefore, we, for instance, based our analysis of the appropriate number of the cointegrating vectors on graphical and estimated coefficients analysis as well as on formal testing.

The small sample correction to the critical values of the test statistics by Reimers (1992, 222) was already introduced in section 5.4. A comparison study of the Johansen procedure in a bivariate model and the Engle-Granger method was conducted by Gonzalo (1990 as quoted in Banerjee et al. 1993, 285) using the Monte Carlo technique. The Johansen ML procedure did well in the study and Gonzalo found that it repeatedly has the smallest mean-squared error among a selection of parameter values of interest to empirical research. He also makes a point that the effects of non-normal errors seem negligible, which, in our case, is a good piece of news. Cheung and Lai (1993, 320) make a similar finding and especially that excessive kurtosis, which is the case with our exchange rate data, is not very influential for inference on cointegration. They also made a similar conclusion as in the above mentioned Johansen (1989, as quoted in Banerjee et al. 1993, 277) study concerning the maximum lag length of the VAR. The tests seem to be biased toward finding cointegration more often than what asymptotic theory suggests and the bias gets larger as the dimension of the system of the lag length increase. Reimers (1992) supports this. Therefore he recommends first to estimate the

[37] 'Near'- or 'borderline'-integrated processes are such that they have a root close to but not on the unit circle (Banerjee et al. 1993, 95).

lag order by model misspecification tests and only after that to specify the cointegration rank.

5.4.3. Testing the robustness of the Johansen ML cointegration tests

In order to assess the robustness of the obtained results, we conducted the cointegration tests on different values of the maximum lag length, k. In the FIM/DEM case the results for k = 5, do not give very different results from the original analysis with k = 4 and one cointegrating vector. The maximum length of six lags, on the other hand, indicates the acceptance of three cointegrating vectors. As to the FIM/USD case, lags one, two, and four all lead to accept two cointegrating vectors, i.e. more than in the original analysis of three lags and one or none cointegrating vectors. The results are reported in Appendices 6 and 7.

The indication, thus, is that the 'correct' number of maximum lags used in the VAR analysis is very important. The results may vary quite significantly and the conclusion may be drawn that the Johansen method indeed seems sensitive to the changes in the number of maximum lags used in the analysis. Therefore it seems odd that for instance in a similar cointegration analysis by Clarida and Taylor (1993), which was introduced in section (3.2.), no emphasis at all is put on the model misspecification tests and the appropriate value of maximum lag length in the VAR.

5.4.4. Parameter restriction tests

The Johansen method allows for testing structural hypotheses on both the cointegration and the adjustment space. The former means linear restrictions on the matrix β of cointegration vectors and the latter on the matrix α adjustment parameters. The tests are first introduced on a general level and the applied, on the appropriate scale, to the data.

Johansen and Juselius (1992, 211-212, 225) introduce three types of hypotheses on the cointegration space:

i) the same linear restrictions on all the cointegrating vectors
ii) some cointegrating vectors are known while others remain unrestricted
iii) a general hypothesis that some cointegrating vectors are restricted by linear restrictions while others remain unrestricted

The hypotheses on the cointegration space are formulated as below:

$$(5.6) \quad \mathcal{H}_{\beta}: \beta = (H_1\varphi_1, H_2\varphi_2, \dots, H_r\varphi_r)$$

where the matrices $H_1, \dots, H_r$ express the linear hypothesis to be tested on each of the r cointegrating relations. (Johansen & Juselius, 1994, 14) The H_i, $i = 1, \dots, r$, matrices are ($n \times s_i$) matrices and the φ_i are parameter matrices in dimension ($s_i \times r$), where n is the number of variables in the cointegrating vector, r is the cointegration rank and, at the same time, the number of cointegrating vectors, and s is the number of non-restricted variables, so, that s = n - (the number of restrictions imposed).

The hypotheses introduced above can be stated more specifically as (using the notation in Johansen and Juselius 1992):

i) $\mathcal{H}_4: \beta = H_4\varphi$, which is a special case of (5.6) when $H_1 = H_2 = \ldots = H_r$. The same (n - s) linear restrictions are tested on all the r cointegration vectors, here $r \leq s \leq n$.

ii) $\mathcal{H}_5: \beta = (H_5, \psi)$, where r_1 of the cointegration vectors are known and specified by the (n x r_1) matrix H_5 while the remaining r_2 vectors are unrestricted and specified by the (n x r_2) matrix ψ, $r_1 + r_2 = r$.

iii) $\mathcal{H}_6: \beta = (H_6\varphi, \psi)$, which is a combination of the above two hypotheses. Here some restrictions (φ) are put on the r_1 cointegration vectors, H_6 is a (n x s) matrix, φ is a (s x r_1) matrix, while the r_2 vectors remain unrestricted in a (n x r_2) matrix of ψ, $r_1 \leq s \leq n$, $r_1 + r_2 = r$. The above two hypotheses are indeed special cases of this hypothesis because with the appropriate choice of r_1, r_2, and s they can be obtained from $\mathcal{H}_6$. If $r_2 = 0$, the hypothesis reduces to $\mathcal{H}_4$, and if $r_1 = s$, i.e. r_1 cointegration vectors are known, $\mathcal{H}_6$ reduces to $\mathcal{H}_5$.

Restrictions on the adjustment space, i.e. on matrix α can be expressed as:

(5.7) $\mathcal{H}_7: \alpha = A\psi$

where α is a (n x r), A is a (n x s) matrix and ψ a (s x r) matrix. (Johansen & Juselius 1990, 199) Usually the restricted α is tested in respect of excluding one or more cointegrating vectors from the system X_t. Formally this is called testing for weak exogeneity on ΔX_{it} for α and β: if for some i $\alpha_{i.} = 0$, then ΔX_{it} is weakly exogenous for α and β. This means that when estimating the parameters of the model (equation 4.6), i.e. Γ_i, Π, α and, β, no information is lost from not modelling the determinants of ΔX_{it}, which is then weakly exogenous to the system. The variable ΔX_{it} is thus an explanatory variable in the system, but the X_{it} itself is not modelled. (Johansen & Juselius 1992, 224 and Harris 1995, 98)

The likelihood ratio tests statistics for the above hypotheses are obtained by finding the value of the maximised likelihood function with and without the restrictions implied by the respective hypothesis and then forming the test statistic. The tests on the above hypothesis are all conditional on the rank maintained in the original cointegration analysis, they are all stationary and the likelihood ratio statistics are asymptotically χ^2 distributed with degrees of freedom (n-s)r for $\mathcal{H}_4$, (n-r)r_1 for $\mathcal{H}_5$, (n-s-r_2)r_1 for $\mathcal{H}_6$, and (n-s)r for $\mathcal{H}_7$. (Johansen & Juselius, 1992, 226-227, 231, 234 and Johansen & Juselius 1990, 200 and Doornik & Hendry 1994, 225)

For our purposes the parameter restriction tests in the form of hypotheses $\mathcal{H}_4$ and $\mathcal{H}_7$ are tested on our model for FIM/DEM and FIM/USD. Since the rank in both the cases has been assessed as being no more than one, r = 1, our analysis of parameter restrictions on β is that of the hypothesis $\mathcal{H}_4$ introduced above.

The likelihood ratio test statistic for the $\mathcal{H}_4$ case against the null $\mathcal{H}_1 : \Pi = \alpha\beta'$ is the following:

$$(5.9) \quad -2\ln Q(\mathcal{H}_4|\mathcal{H}_1) = T\sum_{i=1}^{r} \ln\left\{\psi\left(1-\tilde{\lambda}_i\right)/\left(1-\hat{\lambda}_i\right)\right\}$$

which is then χ^2 distributed with (n-s)r degrees of freedom. (Johansen & Juselius 1990, 194) The $\tilde{\lambda}_1 > \ldots > \tilde{\lambda}_r$ eigenvalues are the restricted eigenvalues and $\hat{\lambda}_1 > \ldots > \hat{\lambda}_r$ are the 'original' eigenvalues obtained from equation (4.18) (cf. equations 4.20 and 4.21).

Given the results of the original cointegration tests, an interesting relation seems to hold between the spot rate and the one-month forward rate It seems as if the rates were homogeneous. Therefore we test for both the time series if $\beta_{11} = -\beta_{21}$ and in addition for the FIM/DEM series if the three-month and six-month forward rates are identical, i.e. $\beta_{31} = \beta_{41}$. Thus, the following restrictions on the cointegrating vector, β, are imposed:

(5.8) $\mathcal{H}_i : \beta = H_i \phi, i = 1,2$

where the matrices H_i are (n x s) and s = n - (the number of restrictions), which here equals one. Note, that in the number of variables, n, the constant is the fifth variable. The corresponding restriction matrices H_i, i = 1, 2 are structured below. The β matrix is of the form (n x r) and since r = 1 in our analysis the β is an n-dimensional vector. The test results on all the following tests are reported in Appendices 8 and 9.

$$H_1 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ -1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \qquad H_2 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$$

The results are very clear: for both the time series the hypothesis of homogeneity between spot and one-month forward rate is accepted at 76 percent probability for FIM/DEM and at 59 percent probability for FIM/USD. The second hypothesis that the three-month and six-month forward rates are identical for FIM/DEM is rejected at one percent significance level.

The parameter restriction tests on the adjustment matrix, α, are conducted in a similar manner. The likelihood ratio test statistic for the $\mathcal{H}_7$ case against the null $\mathcal{H}_1 : \Pi = \alpha\beta'$ is of the following form:

$$(5.10)\quad -2\ln Q(\mathcal{H}_7|\mathcal{H}_1) = T\sum_{i=1}^{r}\ln\{(1-\tilde{\lambda}_i)/(1-\hat{\lambda}_i)\}$$

which is then χ^2 distributed with degrees of freedom (n-s)r. (Johansen & Juselius 1992, 224) (cf. equation 5.10)

In the cointegration tests the estimated coefficients for the adjustment matrix were very small. Therefore we first test if they are effectively zero, one at a time, excluding the spot rate. It may noted that they cannot be zero all at the same time, or there would be no endogenous variables left in the system to build the model for.

The hypotheses are:

$$\mathcal{H}_i : \alpha = A_i\psi,\ i = 3, 4, 5$$

where the A_i are presented as (n x s) vectors which correspond to $\alpha_{11} = \alpha_{13} = \alpha_{14}$, $\alpha_{11} = \alpha_{12} = \alpha_{14}$, $\alpha_{11} = \alpha_{12} = \alpha_{13}$, respectively:

$$A_3 = \begin{bmatrix} 1 \\ 0 \\ 1 \\ 1 \end{bmatrix} \qquad A_4 = \begin{bmatrix} 1 \\ 1 \\ 0 \\ 1 \end{bmatrix} \qquad A_5 = \begin{bmatrix} 1 \\ 1 \\ 1 \\ 0 \end{bmatrix}$$

The restrictions were rejected at one percent significance level in every case. It seems that each forward rate is important and cannot be left out of the analysis. Next we test whether the spot rate alone could be included in the short run analysis, and whether each forward rate at a time could accompany it, i.e. hypotheses

$$(5.11) \quad \mathcal{H}_i : \alpha = A_i \psi, \; i = 6, 7, 8, 9$$

and $\alpha_{12} = \alpha_{13} = \alpha_{14} = 0$, $\alpha_{11} = \alpha_{12}$, $\alpha_{11} = \alpha_{13}$, $\alpha_{11} = \alpha_{14}$ with corresponding restriction vectors (A_i, i = 7, 8, 9 show the first columns only, since that specifically includes the restriction):

$$A_6 = \begin{bmatrix} 1 \\ 0 \\ 0 \\ 0 \end{bmatrix} \qquad A_7 = \begin{bmatrix} 1 \\ 1 \\ 0 \\ 0 \end{bmatrix} \qquad A_8 = \begin{bmatrix} 1 \\ 0 \\ 1 \\ 0 \end{bmatrix} \qquad A_9 = \begin{bmatrix} 1 \\ 0 \\ 0 \\ 1 \end{bmatrix}$$

The restrictions are rejected at one percent significance level in every case. Then, because the results for the acceptance of $\beta_{11} = -\beta_{21}$ were so encouraging, we test the restrictions A_i, $i = 6, 7, 8, 9$, under the restriction that $\beta_{11} = -\beta_{21}$.The results are the same as for the adjustment parameter restrictions alone: rejection of the restrictions at one percent significance level.

The implication of the parameter restriction test results is thus that the data is best modelled with a vector error correction model where the error correction term does not include all the four variables of the system X_t but only the spot and one-month forward rates. Since all the restrictions on the adjustment matrix α were rejected, none of the variables is weakly exogenous to the system and we are able to model both their short and long run behaviour.

It may be noted that when $r > 1$ the uniqueness of the cointegrating vectors is not selfevident. Then restrictions must be imposed to obtain the unique vectors within the unique cointegration space, which is given by the Johansen procedure. This is followed by testing whether the columns of the cointegration matrix β are identified. (Johansen & Juselius 1994) When $r = 1$, on the other hand, the cointegration space is uniquely defined by a single vector, which is the case in our study.

6. CONCLUSION

The purpose of the study was to find out whether the forward exchange rates contain information in respect to the spot exchange rates. The starting point of the study was the unbiased efficient expectations hypothesis (UEE) which states, with the assumptions of interest rate parity and rational expectations, that the n-period forward rate at time t is the unbiased predictor of the spot rate at time t+n, given the information set and a possible risk premium known at time t.

The data was chosen to represent the value of the Finnish markka after its floating on 8 September, 1992 against the US Dollar and the German Mark. The forward rates under study were those of the one-, three-, and six-months of maturity. The aim was to establish a long run equilibrium relationship between the four rates, i.e. the spot rate and the three forward rates, by using cointegration methodology. The first step was thus to maintain the non-stationarity of the exchange rates. This was done by Dickey Fuller unit root tests, the results of which confirmed the series to be non-stationary in first differences, which then filled the back-ground assumptions of the cointegration methodology. Because we were dealing with more than two variables, we chose to investigate cointegration with the Johansen maximum likelihood methodology, instead of that of Engle and Granger. For this purpose we first had to assess that the data fulfil the basic assumption of the Johansen procedure: when modelled as a vector error correction model (VECM) the maximum lag length in the VECM must be chosen so that the residuals are normally, independently, and identically distributed. The conducted model misspecification tests report no autocorrelation when the maximum lag length for the FIM/DEM data is four and for the FIM/USD data three. The cointegration tests then use these maximum lag lengths when determining the number of cointegrating vectors in the time series. The normality assumption was failed by the data, which showed signs of kurtosis. Since kurtostic data has been determined less serious than skewed data in the cointegration analysis, we continued the analysis.

For both the time series we were able to identify one cointegrating vector, i.e. long run equilibrium relation, the existence of which is less self-evident for the FIM/USD data than for the FIM/DEM data. The reason could be the tendency of the Dollar exchange rate to react to almost any piece of news reaching the market. Therefore the 'prediction periods' used in the study, one, three, and six months, may be too long.

To see the effect of the forward rates of different time horizons, we studied the estimated cointegrating and adjustment parameters and tested restrictions on them. According to the tests no forward rate is exogenous to the system and for both the data sets the spot rate seems to be identical to the one-month forward rate, them being the only variables in the equilibrium relation. This particular finding indeed is in accordance with the UEE hypothesis. But since the exogeneity of the forward rates was rejected in every case, the strongest result of the study is that all the three forward rates are important in determining the cointegrating relation, the existence of which implies that the forward exchange rates do contain information about the future spot rate, even if they are not unbiased estimators of the spot rate. It must be kept in mind that the results apply only given the several decisions made about the testing procedure in this particular study.

As a suggestion for further research, it would be interesting to conduct the study with a longer observation period and maybe to include the twelve-month forward exchange rate in the model to see if it is useful to the system or not. Other tests induced by the cointegration methodology, like the one introduced in section 3.2. by Baillie and

Bollerslev (1989) to investigate the interdependence of two spot rates, would also provide an interesting research area for the Finnish exchange rate data.

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ABBREVIATIONS EXPLAINED

Currencies:

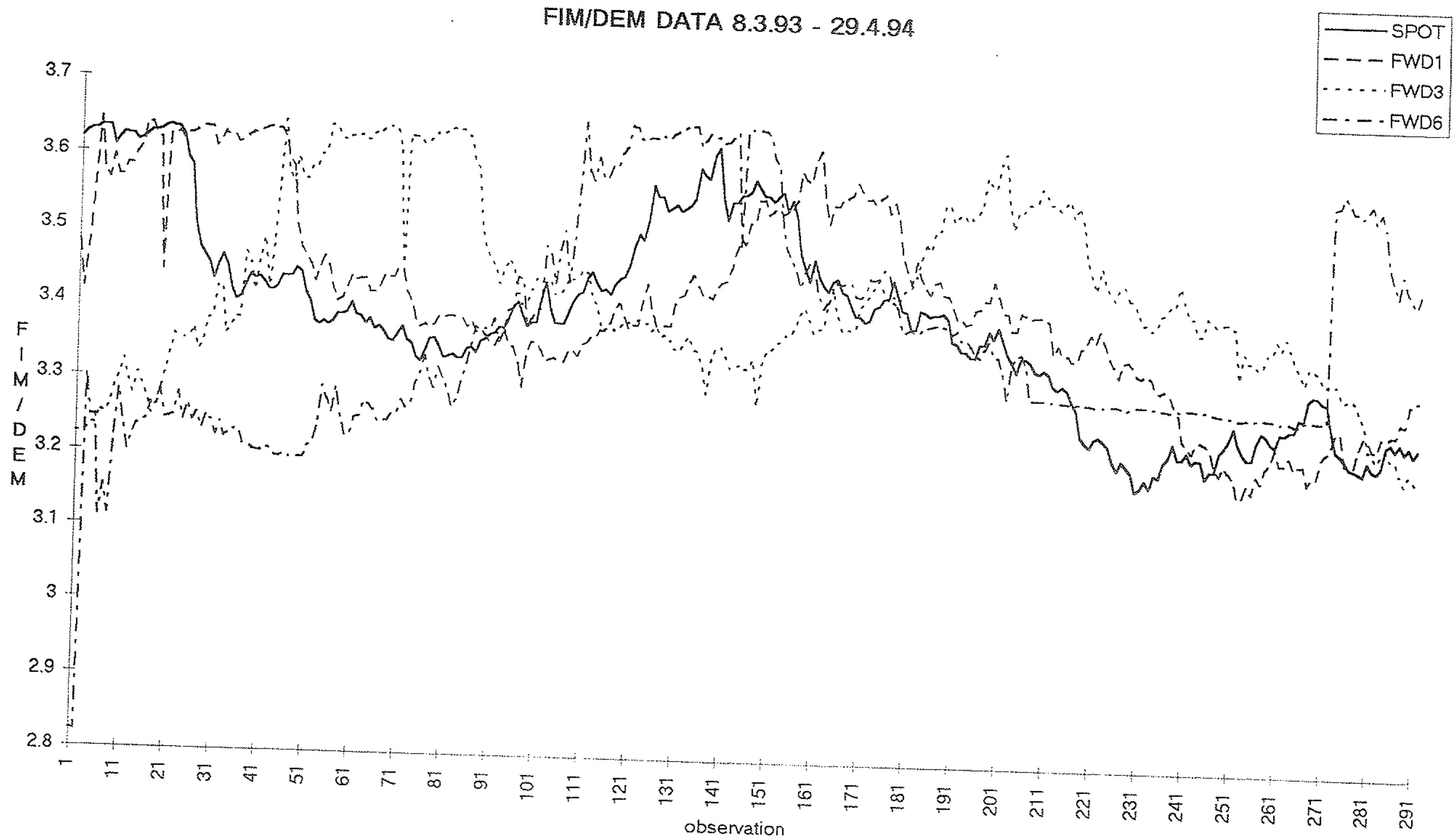
ATS	Austrian Schilling
BEF	Belgian Franc
CAD	Canadian Dollar
DEM	German Mark
DKK	Danish Krone
ESP	Spanish Peseta
FIM	Finnish Markka
FRF	French Franc
GBP	British Sterling
HKD	Hong Kong Dollar
ITL	Italian Lira
NLG	Dutch Gilder
NOK	Norwegian Krone
NZD	New Zealand Dollar
SEK	Swedish Krona
SFR	Swiss Franc
SGD	Singapore Dollar
USD	US Dollar
JPY	Japanese Yen

Mathematical Terms:

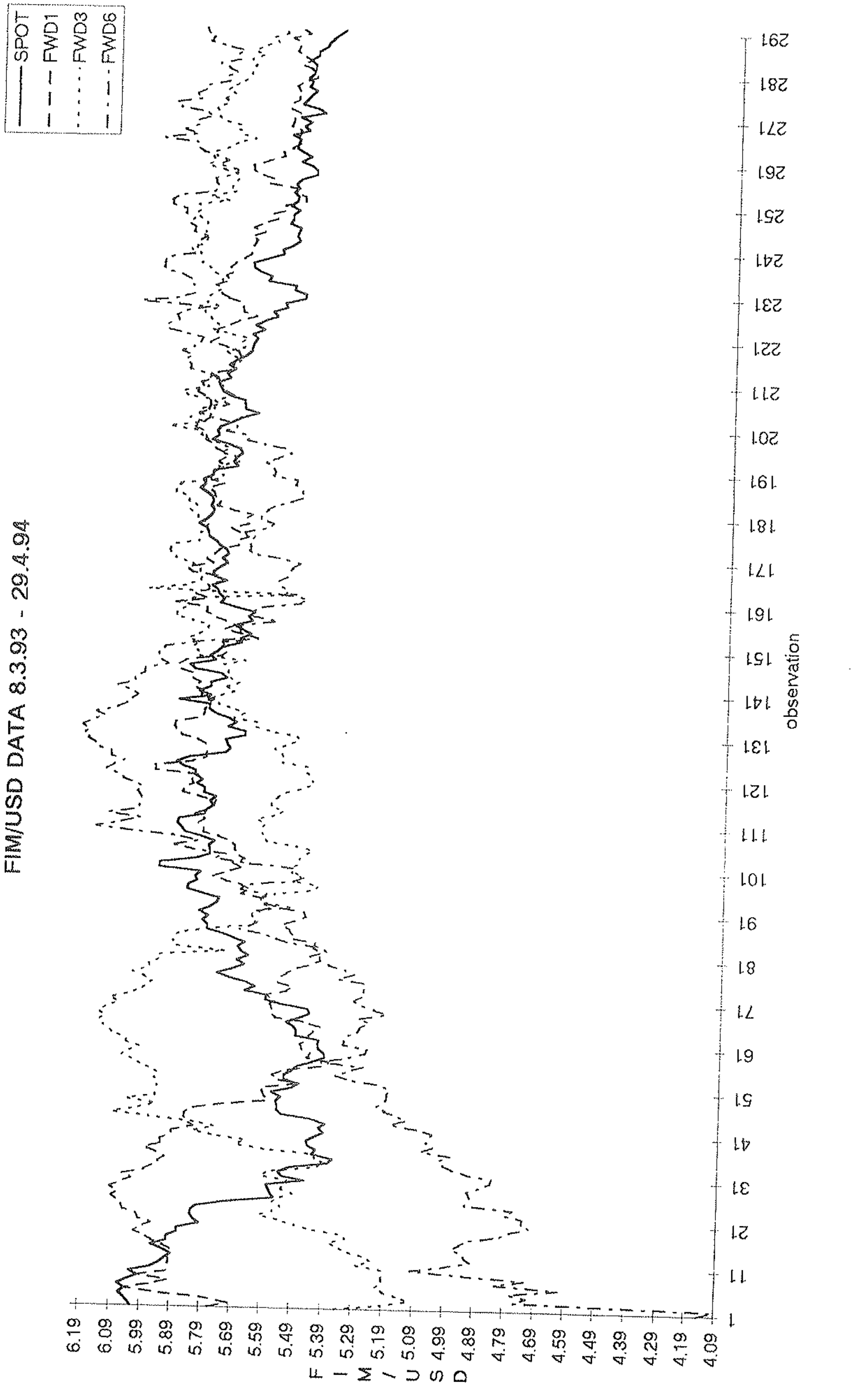
ECM	Error Correction Model
FIML	Full Information Maximum Likelihood
ML	Maximum Likelihood
OLS	Ordinary Least Squares
VECM	Vector Error Correction Model

Variables Used in the Analysis:

SPOT	spot exchange rate
FWD1	one-month forward exchange rate
FWD3	three-month forward exchange rate
FWD6	six-month forward exchange rate


FIM/DEM DATA 8.3.93 - 29.4.94
SPOT
FWD1
FWD3
FWD6
FIM/DEM
3.7
3.6
3.5
3.4
3.3
3.2
3.1
3
2.9
2.8
1
11
21
31
41
51
61
71
81
91
101
111
121
131
141
151
161
171
181
191
201
211
221
231
241
251
261
271
281
291
observation

APPENDIX 3



FIM/USD DATA 8.3.93 - 29.4.94
SPOT
FWD1
FWD3
FWD6
FIM/USD
6.19
6.09
5.99
5.89
5.79
5.69
5.59
5.49
5.39
5.29
5.19
5.09
4.99
4.89
4.79
4.69
4.59
4.49
4.39
4.29
4.19
4.09
1
11
21
31
41
51
61
71
81
91
101
111
121
131
141
151
161
171
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261
271
281
291
observation

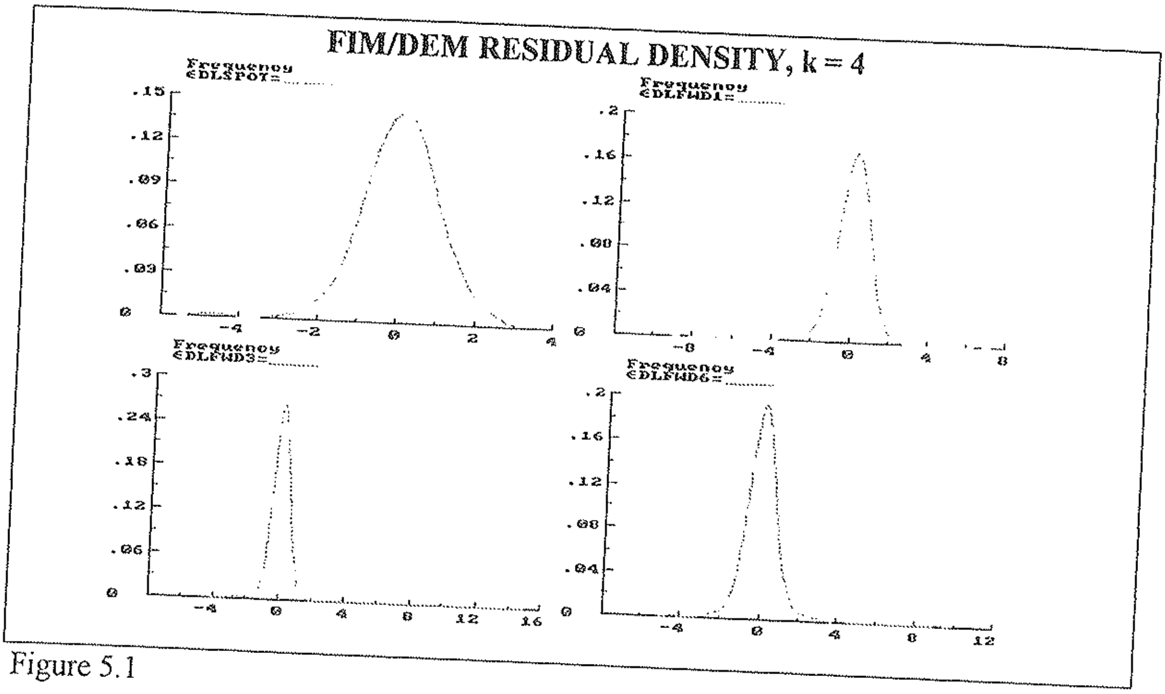

Figure 5.1

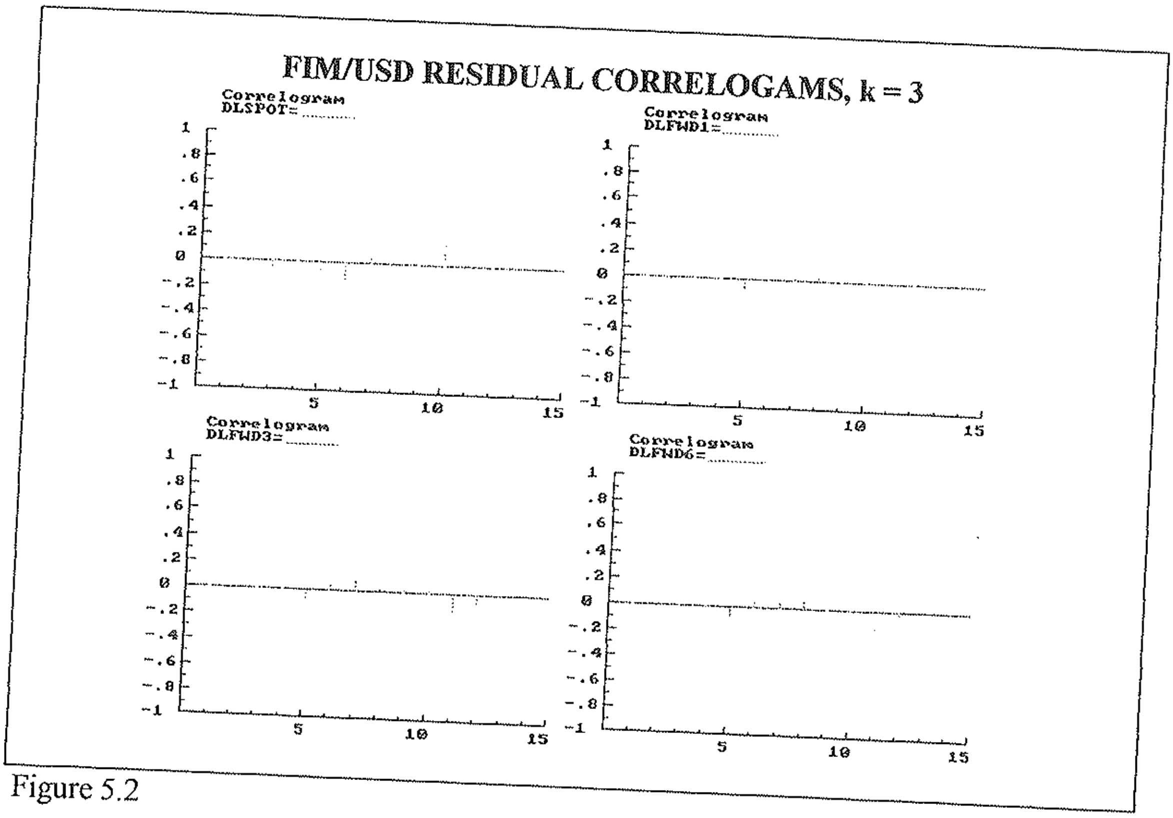

Figure 5.2

COINTEGRATING RELATIONS $\hat{\beta}_i' X_t$, i = 1, 2, 3, 4 FOR FIM/DEM

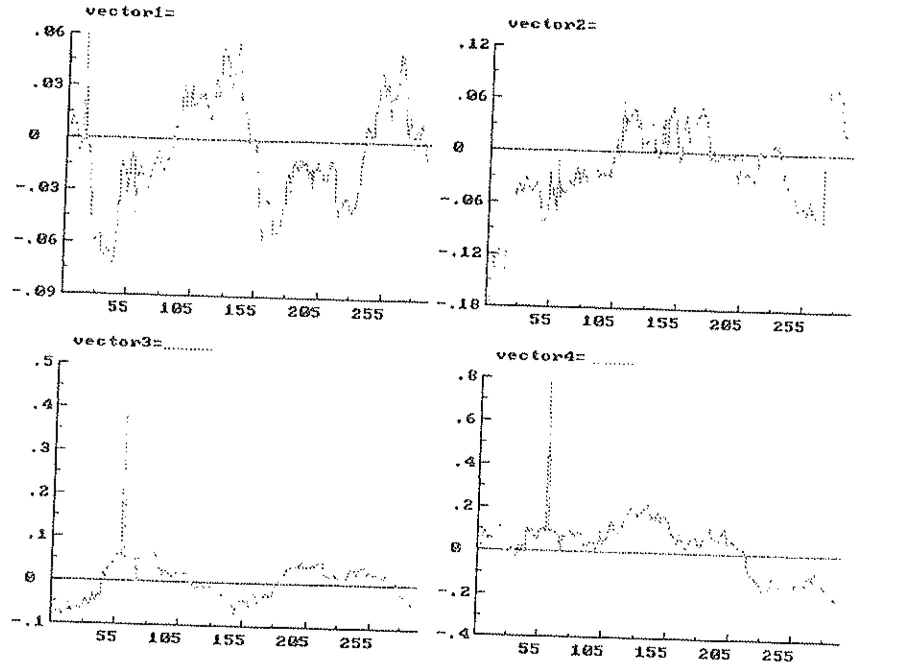

Figure 5.3

COINTEGRATING RELATIONS $\hat{\beta}_i' X_t$, i = 1, 2, 3, 4 FOR FIM/USD

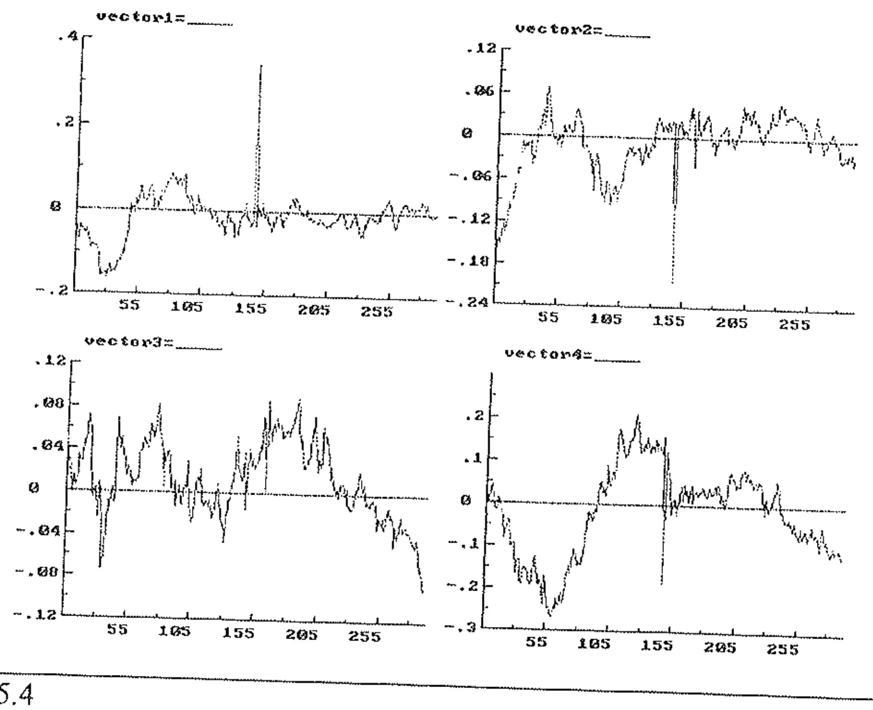

Figure 5.4

TESTS OF THE COINTEGRATION RANK OF THE FIM/DEM DATA, k=5

H_0: rank = r	$\hat{\lambda}_i$	$-T\ln(1-\hat{\lambda}_{r+1})$	$\hat{\lambda}_{max}(0.95)$	$\hat{\lambda}_{max}(T-nk)$
r = 0		32.34**	28.1	30.07**
r ≤ 1	0.107257	19	22.0	17.67
r ≤ 2	0.0645081	14.51	15.7	13.49
r ≤ 3	0.0496402	3.071	9.2	2.856

H_0: rank = r	$\hat{\lambda}_i$	$-T\sum\ln(1-\hat{\lambda}_i)$	$\hat{\lambda}_{trace}(0.95)$	$\hat{\lambda}_{trace}(T-nk)$
r = 0		68.92***	53.1	64.09***
r ≤ 1	0.107257	36.59**	34.9	34.9
r ≤ 2	0.0645081	17.58	20.0	16.35
r ≤ 3	0.0496402	3.071	9.2	2.856

** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 5% significance level.

*** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 1% significance level.

TESTS OF THE COINTEGRATION RANK OF THE FIM/DEM DATA, k=6

H_0: rank = r	$\hat{\lambda}_i$	$-T\ln(1-\hat{\lambda}_{r+1})$	$\hat{\lambda}_{max}(0.95)$	$\hat{\lambda}_{max}(T-nk)$
r = 0		28.44**	28.1	26.05
r ≤ 1	0.0949832	21.02**	22.0	19.25
r ≤ 2	0.0710928	16.85**	15.7	15.43
r ≤ 3	0.0574064	2.225	9.2	2.037

H_0: rank = r	$\hat{\lambda}_i$	$-T\sum\ln(1-\hat{\lambda}_i)$	$\hat{\lambda}_{trace}(0.95)$	$\hat{\lambda}_{trace}(T-nk)$
r = 0		68.54***	53.1	62.76***
r ≤ 1	0.0949832	40.09**	34.9	36.72**
r ≤ 2	0.0710928	19.07	20.0	17.47
r ≤ 3	0.0574064	2.225	9.2	2.037

** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 5% significance level.

*** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 1% significance level.

TESTS OF THE COINTEGRATION RANK OF THE FIM/USD DATA, k=1

H_0: rank = r	$\hat{\lambda}_i$	$-T\ln(1-\hat{\lambda}_{r+1})$	$\hat{\lambda}_{max}(0.95)$	$\hat{\lambda}_{max}(T-nk)$
r = 0		55.31***	28.1	54.54***
r ≤ 1	0.175295	23.73**	22.0	23.4**
r ≤ 2	0.0793685	8.62	15.7	8.5
r ≤ 3	0.0295885	3.413	9.2	3.365

H_0: rank = r	$\hat{\lambda}_i$	$-T\sum\ln(1-\hat{\lambda}_i)$	$\hat{\lambda}_{trace}(0.95)$	$\hat{\lambda}_{trace}(T-nk)$
r = 0		91.08***	53.1	89.81***
r ≤ 1	0.175295	35.77**	34.9	35.27**
r ≤ 2	0.0793685	12.03	20.0	11.87
r ≤ 3	0.0295885	3.413	9.2	3.365

** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 5% significance level.

*** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 1% significance level.

TESTS OF THE COINTEGRATION RANK OF THE FIM/USD DATA, k=2

H_0: rank = r	$\hat{\lambda}_i$	$-T\ln(1-\hat{\lambda}_{r+1})$	$\hat{\lambda}_{max}(0.95)$	$\hat{\lambda}_{max}(T-nk)$
r = 0		30.72**	28.1	29.86**
r ≤ 1	0.101508	23.49**	22.0	22.83**
r ≤ 2	0.0785851	6.262	15.7	6.087
r ≤ 3	0.0215814	3.68	9.2	3.577

H_0: rank = r	$\hat{\lambda}_i$	$-T\sum\ln(1-\hat{\lambda}_i)$	$\hat{\lambda}_{trace}(0.95)$	$\hat{\lambda}_{trace}(T-nk)$
r = 0		64.15***	53.1	62.36***
r ≤ 1	0.101508	33.43	34.9	32.5
r ≤ 2	0.0785851	9.941	20.0	9.664
r ≤ 3	0.0215814	3.68	9.2	3.577

** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 5% significance level.

*** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 1% significance level.

TESTS OF THE COINTEGRATION RANK OF THE FIM/USD DATA, k=4

H_0: rank = r	$\hat{\lambda}_i$	$-T\ln(1-\hat{\lambda}_{r+1})$	$\hat{\lambda}_{max}(0.95)$	$\hat{\lambda}_{max}(T\text{-}nk)$
$r = 0$		29.54**	28.1	27.89
$r \leq 1$	0.0978124	23.15**	22.0	21.86
$r \leq 2$	0.0775011	5.113	15.7	4.828
$r \leq 3$	0.0176567	3.242	9.2	3.061

H_0: rank = r	$\hat{\lambda}_i$	$-T\sum\ln(1-\hat{\lambda}_i)$	$\hat{\lambda}_{trace}(0.95)$	$\hat{\lambda}_{trace}(T\text{-}nk)$
$r = 0$		61.05***	53.1	57.65**
$r \leq 1$	0.0978124	31.51	34.9	29.75
$r \leq 2$	0.0775011	8.355	20.0	7.889
$r \leq 3$	0.0176567	3.242	9.2	3.061

** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 5% significance level.

*** indicates rejection of the $\mathcal{H}_1$ in $\hat{\lambda}_{max}$ and $\mathcal{H}_0$ in $\hat{\lambda}_{trace}$ at 1% significance level.

PARAMETER RESTRICTION TESTS ON FIM/DEM DATA

	H_3	H_4	H_5
α-restrictions	$\alpha_{12}=0$	$\alpha_{13}=0$	$\alpha_{14}=0$
$\chi^2(df)$	df=3	df=3	df=3
LR-test	22.12***	10.62**	12.136***
β_{11}	1.000	1.000	1.000
β_{21}	-34.18	-1.645	-1.922
β_{31}	-4.58	-0.161	-0.661
β_{41}	-0.439	0.031	0.541
μ	46.58	0.951	1.270
α_{11}	0.0004	0.019	0.012
α_{12}	0.000	-0.019	0.012
α_{13}	0.0004	-0.000	0.012
α_{14}	0.0004	0.019	0.000

H_{ij}, i = 1, . . . , 9, j = 6, . . . , 9

** indicates rejection of the H_{ij} at 5% significance level
*** indicates rejection of the H_{ij} at 1% significance level

	H_1	H_2	H_6	H_7	H_8	H_9	$H_{1,6}$	$H_{1,7}$	$H_{1,8}$	$H_{1,9}$
β-restrictions	$\beta_{11}=-\beta_{21}$	$\beta_{31}=\beta_{41}$					$\beta_{11}=-\beta_{21}$	$\beta_{11}=-\beta_{21}$	$\beta_{11}=-\beta_{21}$	$\beta_{11}=-\beta_{21}$
α-restrictions			$\alpha_{12}=\alpha_{13}$ $\alpha_{12}=\alpha_{14}=0$	$\alpha_{11}=\alpha_{12}$	$\alpha_{11}=\alpha_{13}$	$\alpha_{11}=\alpha_{14}$	$\alpha_{12}=\alpha_{13}$ $\alpha_{12}=\alpha_{14}=0$	$\alpha_{11}=\alpha_{12}$	$\alpha_{11}=\alpha_{13}$	$\alpha_{11}=\alpha_{14}$
$\chi^2(df)$	df=4	df=4	df=3	df=3	df=3	df=3	df=7	df=7	df=7	df=7
LR-test	1.868(0.76)	25.432***	19.304***	10.811**	21.105***	20.33***	27.92***	17.28***	28.09***	26.78***
β_{11}	1.000	0.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
β_{21}	-1.000	0.000	0.109	-2.020	-0.100	-16.56	-1.000	-1.000	-1.000	-1.000
β_{31}	0.000	1.000	0.012	-0.4134	0.254	0.994	0.000	0.000	0.000	0.000
β_{41}	0.000	1.000	-0.762	0.6213	-0.609	1.204	0.000	0.000	0.000	0.000
μ	0.000	0.000	-0.423	0.9878	-0.6534	16.27	0.000	0.000	0.000	0.000
α_{11}	0.005	-0.0001	-0.016	0.0121	-0.017	0.001	0.006	0.029	0.004	0.011
α_{12}	0.061	-0.0001	0.000	0.0121	0.000	0.000	0.000	0.029	0.000	0.000
α_{13}	-0.041	-0.0000	0.000	0.000	-0.017	0.000	0.000	0.000	0.004	0.000
α_{14}	0.023	0.0002	0.000	0.000	0.000	0.001	0.000	0.000	0.000	0.011

PARAMETER RESTRICTION TESTS ON FIM/USD DATA

α-restrictions	H_3 $\alpha_{12}=0$	H_4 $\alpha_{13}=0$	H_5 $\alpha_{14}=0$
$\chi^2(df)$	df=3	df=3	df=3
LR-test	23.60**	23.69**	24.82**
β_{11}	1.000	1.000	1.000
β_{21}	-0.271	16.96	0.545
β_{31}	0.318	0.749	0.616
β_{41}	0:151	2.714	-0.011
μ	-2.081	-37.27	-3.737
α_{11}	-0.025	-0.002	-0.022
α_{12}	0.000	-0.002	-0.022
α_{13}	-0.025	0.000	-0.022
α_{14}	-0.025	-0.002	0.000

H_{ij}, i = 1, . . . , 9, j = 6, . . . , 9

** indicates rejection of the H_{ij} at 5% significance level

*** indicates rejection of the H_{ij} at 1% significance level

β-restrictions / α-restrictions	H_1 $\beta_{11}=-\beta_{21}$	H_2 $\beta_{31}=\beta_{41}$	H_6 $\alpha_{12}=\alpha_{13}$ $\alpha_{12}=\alpha_{14}=0$	H_7 $\alpha_{11}=\alpha_{12}$	H_8 $\alpha_{11}=\alpha_{13}$	H_9 $\alpha_{11}=\alpha_{14}$	$H_{1,6}$ $\beta_{11}=-\beta_{21}$ $\alpha_{12}=\alpha_{13}$ $\alpha_{12}=\alpha_{14}=0$	$H_{1,7}$ $\beta_{11}=-\beta_{21}$ $\alpha_{11}=\alpha_{12}$	$H_{1,8}$ $\beta_{11}=-\beta_{21}$ $\alpha_{11}=\alpha_{13}$	$H_{1,9}$ $\beta_{11}=-\beta_{21}$ $\alpha_{11}=\alpha_{14}$
$\chi^2(df)$	df=4		df=3	df=3	df=3	df=3	df=7	df=7	df=7	df=7
LR-test	2.795(0.59)		20.044**	19.048**	23.189**	25.33**	29.22**	27.156**	25.993**	29.035**
β_{11}	1.000		1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
β_{21}	-1.000		0.245	7.145	-0.319	-0.436	-1.000	-1.000	-1.000	-1.000
β_{31}	0.000		-0.757	-0.763	0.176	-0.551	0.000	0.000	0.000	0.000
β_{41}	0.000		-0.161	-0.424	-0.034	0.297	0.000	0.000	0.000	0.000
μ	0.000		-0.521	-12.05	-1.428	-2.046	.000	0.000	0.000	0.000
α_{11}	-0.004		-0.022	-0.005	-0.037	-0.013	-0.003	0.019	-0.021	-0.005
α_{12}	0.138		0.000	-0.005	0.000	0.000	0.000	0.019	0.000	0.000
α_{13}	-0.055		0.000	0.000	-0.037	0.000	0.000	0.000	-0.021	0.000
α_{14}	-0.005		0.000	0.000	0.000	-0.013	0.000	0.000	0.000	-0.005

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