STUDIECENTRUM VOOR ECONOMISCH EN SOCIAAL ONDERZOEK

VAKGROEP MACRO-ECONOMIE

ESTIMATING AND FORECASTING EXCHANGE RATES BY MEANS OF PPP AND UIP

An application on the Belgian Franc and the Dutch Guilder vis à vis the DM and US Dollar

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ABSTRACT

Recently, the interest for international parity-conditions has increased, due to the integration of the world capital and goods markets. As such, the Purchasing Power Parity(PPP)- and the Uncovered Interest Parity(UIP)-hypothesis seem to be important. This is the issue of the underlying paper : the PPP- and the UIP-hypotheses are examined for four exchange rates, namely the Belgian Franc and the Dutch Guilder vis à vis the US Dollar and the DM for the sampling periods (namely, weekly data from wedn. 2 January 1980 till wedn. 28 November 1990 and monthly data from January 1980 till November 1990).

In a first section, the PPP-hypothesis is examined by Dickey-Fuller unit root tests on the stationarity of the exchange rates and relative consumer price indices and by co-integration tests on the long-run validity.

In a second section, the UIP-hypothesis is tested in an analogous way to the PPP-hypothesis.

In a third section, it is tried to come to an alternative explanation of the exchange rates. On the basis of the outcomes of causality tests between the exchange rates on the one side and the relative consumer price indices and the interest rate differentials on the other hand, transfer functions to estimate an equation for each (logarithmic) exchange rate are set up and are used to generate ex-ante forecasts for the above mentioned exchange rates.

INTRODUCTION

This paper aims at constructing a model for the exchange rate (returns) for the Belgian Franc and the Dutch Guilder vis à vis the US Dollar and the DM. Many studies conclude that the exchange rate (returns) follow a random walk process. For example, Meese and Singleton (1982) found random walk patterns in the weekly spot and forward exchange rates of the Swiss Franc, the DM and the Canadian Dollar vis à vis the US Dollar during 1976-81 and, more recently, Diebold and Nerlove (1986) came to the same conclusion for the weekly spot rates of the DM, the Japanese Yen and the Canadian Dollar vis à vis the US Dollar during 1973-85.

However, one should be cautious to conclude too quickly that the exchange rate should necessarily follow or did in fact follow a random walk. A first reason is that one has to consider that a random walk requires quite stringent conditions about the time-series process of underlying economic variables, e.g. constant expected returns. In this sense, Takagi (1986) showed that in a formal model, random walk implies the absence of correlation between the exchange rate and the interest differential. Another reason is that the Dickey-Fuller types of unit root tests have low power against borderline stationarity alternatives. On the basis of a Monte Carlo study, Hakkio (1986) showed that four popular types of random walk tests, including the Dickey-Fuller test, have an extremely low rejection rate when the true model follows a stationary process that is close to a random walk.

Recently, Koedijk and Schotman (1990) demonstrated that monthly changes of (the logaritms of) all possible bilateral exchange rates between the United States, the United Kingdom, Germany and Japan for the sampling period - February '77 till July '87 - do not follow a random walk at all. Price differentials between wholesale and consumption prices and interest rates were observed to possess significant impacts on the above mentioned exchange rate returns.

This paper takes the above findings as starting point and performs statistical tests for the Purchasing Power Parity- and the Uncovered Interest Parity-hypotheses for exchange rates. These tests lead to causality tests and to the construction of transfer functions for the above mentioned (logarithmic) exchange rate returns w.r.t. relative prices and interest rate differentials. These transfer functions imply considerable deviations from the random walk hypothesis of the exchange rate returns. Sampling estimates for weekly data from wedn. 2 January 1980 till 28 November 1990 and for monthly data from January 1980 till November 1990 with ex ante forecasts for a period of 8 weeks, 22 weeks and 5 months respectively are presented.

1. The theory of purchasing Power Parity

After a brief introduction, the hypothesis that the elasticity of the exchange rate with respect to the relative consumption price ratio is about unity will be tested in the second paragraph. In the third paragraph the long-run validity of the Purchasing Power Parity(PPP)-hypothesis will be tested using the co-integration methodology proposed by Engle and Granger(1987).

1.1. Introduction

The main idea of the PPP-hypothesis is that exchange rates and national consumption price levels will adjust so as to maintain a given currency's purchasing power across boundaries. This means that under strict PPP-hypothesis the real value of a given currency will be the same in all countries at any moment in time. This can be expressed by the following equation :

$$s_t = \alpha + \beta(p_t - p_t^*) + u_t \qquad (1.1)$$

with :

s_t : the (logarithmic) exchange rate (foreign currency expressed in domestic currency)

p_t : the (logarithm of the) domestic consumer price index

p^*_t : the (logarithm of the) foreign consumer price index

u_t : a white noise error term ($u_t \overset{i.i.d.}{\sim} (0,\sigma^2)$)

Under strict PPP-hypothesis, the following restrictions should be valid : $\alpha=0$; $\beta=1$; $u_t=0$ ($\forall t$). Equation (1.1) corresponds to the absolute version of PPP. The corresponding relative version of PPP is given by :

$$\Delta s_t = \beta\Delta(p_t - p_t^*) + \Delta u_t \tag{1.2}$$

where Δ denotes the first difference operator. The distinction between the absolute and the relative version of the strict PPP-relationship is that in the latter only the percentage change over time of a currency's purchasing power is equalized. In fact, the relative version of the PPP-relationship is a relaxation of the strong hypothesis of the absolute version that $\beta=1$ since due, for example, to tariffs or other distortions, β in eq (1.1) can differ from 1. Nevertheless, if those distortions are constant (in percentage terms), β in equation (1.2) will be constant and can be equal to one, so that the relative version of the PPP holds. The strict PPP-hypothesis, i.e., the restrictions $\alpha=0$ (only equation (1.1)), $\beta=1$, will be tested by an F-test in the next paragraph. Moreover, the long-run validity of the PPP-hypothesis will be tested in a third paragraph, using the co-integration methodology proposed by Engle and Granger(1987).

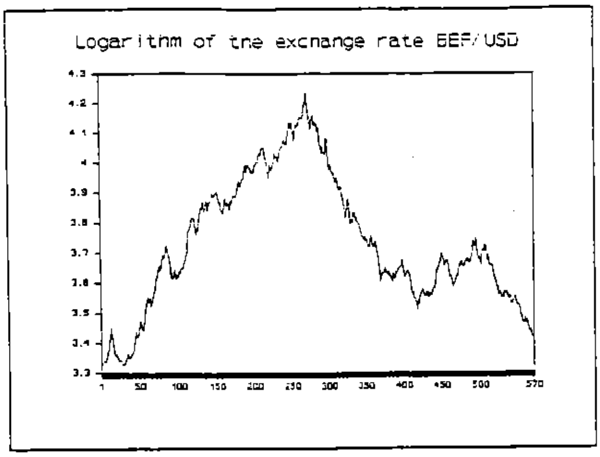
Let us first, before performing these tests, take a brief look at the data. In our analysis, firstly, weekly unadjusted data (wednesday data) from wedn. 2 January 1980 to wedn. 28 November

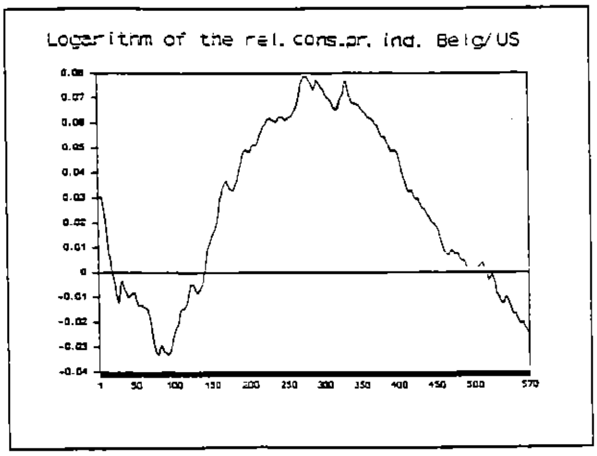
1990 and secondly, monthly data from January 1980 till November 1990 are used. If, for some week, no wednesday datum for the exchange rate was available because of holiday, this datum was generated by interpolation. The PPP-hypothesis is examined for the Belgian Franc and the Dutch Guilder vis à vis the US Dollar and the DM respectively.

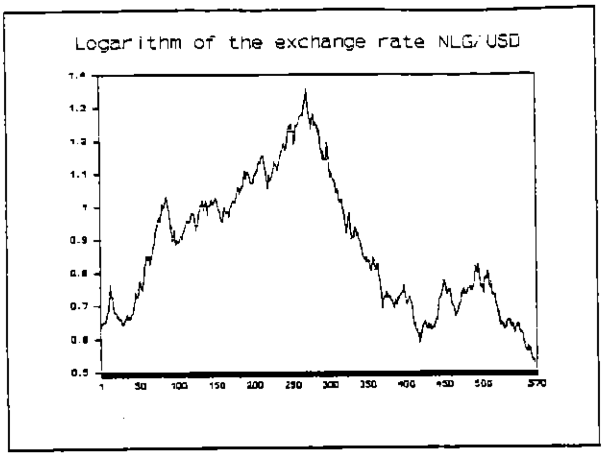
In figures 1.1 to 1.8, the time series of the exchange rates and the relative consumer price indices (all expressed in logarithms) for weekly data are shown. From these figures, we can firstly deduce that the nominal exchange rates fluctuate more than the corresponding relative consumer price indices. This let us suspect the β-coëfficient to be larger than one. This phenomenon - much higher variability in exchange rates than in the relative consumer price indices - was already observed for the period of floating exchange rates from 1975 to 1979. This brought Frenkel(1981) to formulate the hypothesis of the collapse of the Purchasing Power Parity-hypothesis during the seventies.

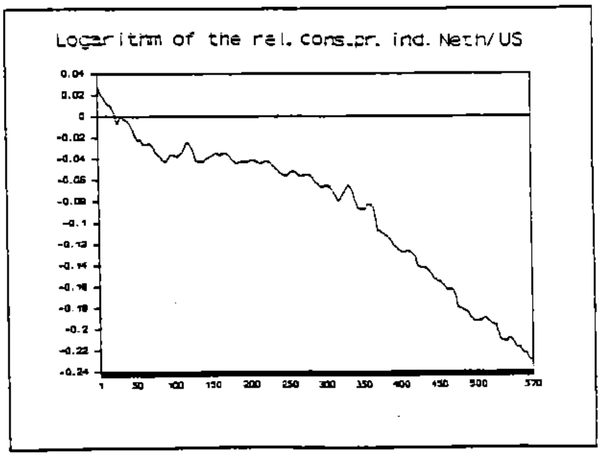
Also other empirical studies on the validity of PPP during the recent floating exchange rate regime as well as during the 1920's confirmed Frenkel's hypothesis that the PPP exist during the 1920's but collapsed during the 1970's. As such, many studies were favourable for the PPP-hypothesis during the 1920's. For example, Frenkel(1978,1980), using traditional techniques, showed that there exists strong evidence supportive to PPP during the 1920's for the bilateral exchange rates involving the US Dollar, the Pound Sterling and the French Franc. To the same conclusion came Taylor and MC Mahon(1988), who examined all bilateral exchange rates

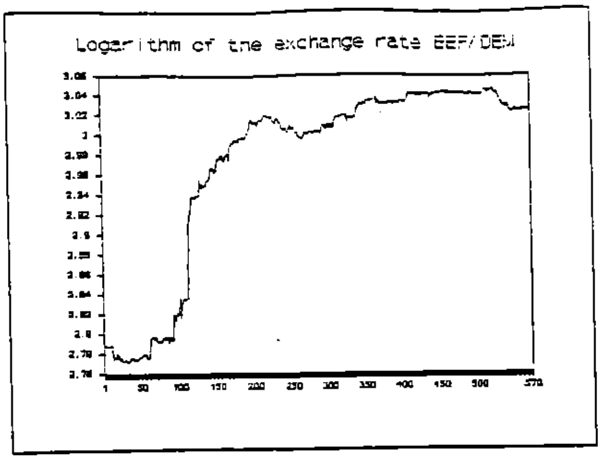
Figures 1.1 - 1.8

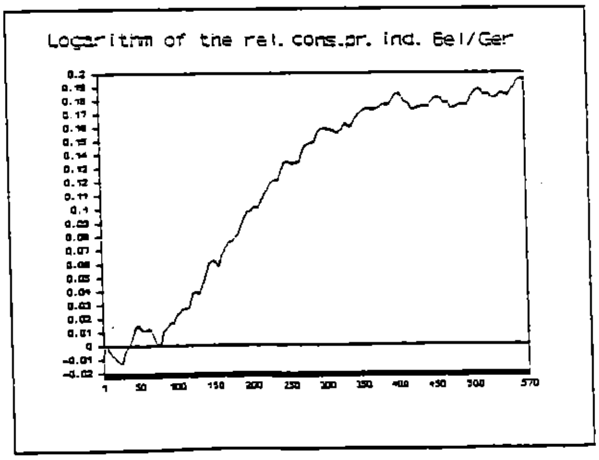


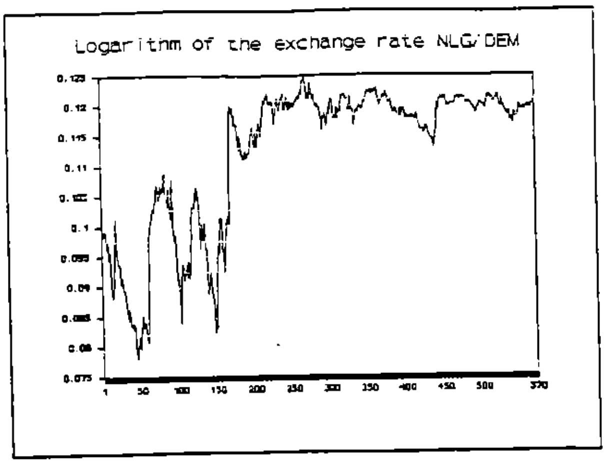


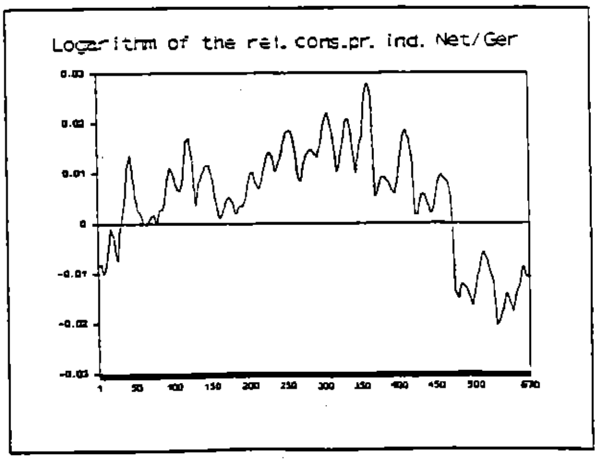












involving all the currencies mentioned above as well as the German mark using the recently developed cointegration techniques. There was one exchange rate for which the PPP did not hold in the long-run, namely the US Dollar-Pound Sterling exchange rate.

In contrast, empirical studies covering the 1970's and 1980's pointed towards the collapse of the PPP-hypothesis (see e.g. Dornbusch 1980[1], Frenkel 1981, Kravis et al. 1975 and Stockman 1978). Studies carried out by e.g. Roll(1979), Piggot and Sweeney(1985) and Adler and Lehmann(1983), in which the time series properties of the (logarithm of the) real exchange rate were examined, concluded that the analysed time series follow a random walk pattern, which implies that the nominal exchange rate and relative consumer price indices do not move together, but tend to drift apart. Consequently, the PPP-hypothesis cannot hold in the long-run.

The collapse of the PPP was explained by Frenkel(1981) by "...the volatile character of the 1970's which witnessed great turbulence in the world economy and large volumes of real shocks like the oil embargo, commodity supply shocks, booms and shortages, shifts in the demands for money and differential productivity growth"(p.162). Consequently, one expects more supportive evidence to PPP for economies where monetary shocks have dominated real shocks. For example, Mussa(1979) argues that PPP performs best for countries

[1] Note that Dornbusch(1976) explained the higher volatility of the exchange rates in comparison with the relative consumer price indices by an initial overshooting of the exchange rates, where he assumed a slow adjustment speed of the prices in the goods markets relative to those of the asset markets.

which have over long periods of time high domestic inflation, high rates of money growth and few supply shocks. Baillie(1990) found more evidence that PPP holds when monetary factors are dominant, in a paper in which he examined long-run PPP during the 1970's and 1980's using the technique of cointegration of economic time series for three high inflation economies in Latin America, namely Argentina, Brazil and Chile.

A second remark w.r.t. the figures is that the figures let us suspect that the time series are non-stationary. In the third paragraph stationarity tests will be performed and the coherency between variables (here between the relative consumer price indices and the exchange rates) will be analysed.

1.2. Testing the hypothesis that the elasticity of the exchange rate with respect to the price ratio is about unity

The aim of this paragraph is to test whether the strict PPP-hypothesis holds. This can be tested by an F-test, which tests the restrictions $\alpha=0$ and $\beta=1$ in equations (1.1) and (1.2).
The outcomes of the estimation of equations (1.1) and (1.2) and of the F-tests are tabulated in table 1.1 of the appendix. The results are little satisfactory : for the relative version of the PPP-relationship the R^2's are very low, while for the absolute version of the PPP-relationship, the DW's are very low, which points to the existence of positive autocorrelation. Moreover, the strict PPP-hypothesis doesn't seem to hold : the outcomes of the F-test lead in every case to the rejection (with a confidence level of 95%) of

the strict PPP-hypothesis with exception of the relative version of the PPP-relationship for the Belgian Franc vis à vis the US Dollar for both weekly and monthly data.

1.3. Stationarity and co-integration between variables

The aim of this paragraph is to analyse whether the PPP-hypothesis holds in the long-run. To test the long-run validity of the PPP-hypothesis the co-integration methodology proposed by Engle and Granger (1987) is used. With the concept of co-integration, Engle and Granger wanted to explain the situation where a particular linear combination of non-stationary variables is itself stationary, which means that a long-run equilibrium between these non-stationary variables exists. The concept of co-integration can be formalized as follows: the components of the vector $x_t := (x_{1t}, x_{2t}, \ldots, x_{kt})'$ are said to be co-integrated of order d,b, denoted $x_t \sim I(d,b)$, if (i) all components of $x_t \sim I(d)$ $\forall i$: 1 to k [2] and (ii) there exists a vector $\alpha (\neq 0)$ so that the linear combination $z_t = \alpha' x_t$ is integrated of order d-b, $z_t \sim I(d-b)$, with b>0 and α called the 'co-integrating vector'. Hence, the definition of co-integration states that a set of time series, all of which become stationary after differencing d times, may have linear combinations which are stationary after differencing b times less. As a consequence, the first thing we have to do, if we want to test whether a co-integrating-relation between the relative prices and the

[2] Note that a series is said to be integrated of order d, denoted $x_t \sim I(d)$, if it has a stationary and invertible ARMA-representation after differencing d times.

corresponding exchange rates exists or not, is to test for the order of integration of the concerning time series. This can be made by computing the following three OLS regressions, defined as:

$$x_t = \rho x_{t-1} + \epsilon_t \qquad (1.3)$$

$$x_t = \alpha + \rho x_{t-1} + \epsilon_t \qquad (1.4)$$

$$x_t = \alpha + \theta t + \rho x_{t-1} + \epsilon_t \qquad (1.5)$$

with x_t the stochastic time series, ϵ_t a white noise error term such that $\epsilon_t \overset{iid}{\sim} N(0,\sigma^2)$ and t=1,2,...T.

In equations (1.3) and (1.4) the hypothesis of a unit root without (equation (1.3)) and with (equation (1.4)) drift, i.e., $H_0^1 : |\rho|=1$ and $H_0^2 : |\rho|=1 ; \alpha=0$ can be tested against the stationary alternative, i.e., $H_1 : |\rho|<1$. Using equation (1.5), we can test the presence of a deterministic time trend in the time series; in equation (1.5) the hypothesis $H_0^3 : |\rho|=1 ; \alpha=0 ; \theta=0$ is tested against the stationary alternative $H_1 : |\rho|<1$.
Virtually, the hypothesis whether $|\rho|=1$ in equations (1.3-5) is tested by the Dickey-Fuller test statistic (see Dickey & Fuller,(1976)), which is calculated as $T(\hat{\rho}-1)$, where T is the number of observations and $\hat{\rho}$ is the OLS-estimator of ρ in equations (1.3-5). The critical values are tabulated in Dickey & Fuller(1976).

Mourik and Winder (1989) stressed that equation (1.5) is preferred above equation (1.4) to test for the order of integration, because, if α is unknown, equation (1.4) can lead to false conclusions. Nevertheless, if a priori information about the value of α is known, the test for stationarity according to equation (1.4) is more powerful than equation (1.5). If α is different from zero the critical values can be generated through simulation, but usually, the true value of α is unknown. This problem can be solved by the introduction of a time trend in equation (1.4). Evans and Savin (1984) show that, if a time trend is incorporated in the regression relationship, the asymptotic distribution of the t-value of the ordinary least squares estimator of ρ is independent of α, so that the introduction of a time trend as regressor in equation (1.5) is recommended.

An alternative way to test for stationarity, is to compute the following three OLS-regressions(see Dickey and Fuller,1976) :

$$\Delta x_t = \phi x_{t-1} + \epsilon_t \tag{1.6}$$

$$\Delta x_t = \alpha + \phi x_{t-1} + \epsilon_t \tag{1.7}$$

$$\Delta x_t = \alpha + \theta t + \phi x_{t-1} + \epsilon_t \tag{1.8}$$

The nullhypothesis that x_t is non-stationary, i.e., $H_0:\phi=0$ is tested then against the stationary alternative $H_1:\phi<0$.

In our analysis, we applied both tests, i.e., the tests according to equations (1.3-5) and equations (1.6-8) respectively on the levels and the first order differences of the exchange rates and relative consumer price indices. Because all tests led to the same conclusion, i.e., all the level series are I(1), we only reported the results of the alternative test (see eq.1.6-1.8) on the first differences (see table 1.2 of the appendix). The subscripts ϵ, μ and τ are used to refer to equations (1.6), (1.7) and (1.8) respectively. In table 1.3 of the appendix, the critical values for the alternative test are tabulated. Note that in table 1.3 only critical values for T=100 and T=200 are tabulated, nevertheless our sample contains 131 monthly and 570 weekly observations. This can be justified by the fact that for large sample sizes the critical values become almost insensitive for relatively small deviations from the sample size considered. Another remark is that the critical values are greater than the conventional critical values 1,96 if α=5% and 2,58 if α=1%. The reason is that due to the fact that the estimated parameter vector is not asymptotically normally distributed under the nullhypothesis, but skew to the left, the conventional values lead too often to a rejection of the nullhypothesis of non-stationarity.

Above, we showed that the relative consumer price indices(p_t) and exchange rates(s_t) series are integrated of order 1. Following the definition of co-integration of Engle and Granger (p.4-5), it is possible that s_t and p_t are co-integrated of order zero, if there exists a constant b, such that

$$c_t = s_t - bp_t$$

which is I(0).

This means that there exists a linear combination between the relative consumption price and exchange rate series which is stationary, such that the linear combination between these two series can be considered as a long-run equilibrium relationship, in the sense that the two non-stationary series do not drift away, but "move together" in the long-run.

To test for co-integration, we use the Augmented Dickey-Fuller regression t-statistic (investigated thoroughly by Engle and Granger, 1987). The Augmented Dickey-Fuller regression t-statistic (ADF-statistic) tests wether the residuals of the co-integrating regression are stationary, whereby the co-integrating regression is given by :

$$y_t = \alpha + \beta x_t + u_t \tag{1.9}$$

and where y_t and x_t are time series and u_t is the residual term. To test whether u_t is stationary the Augmented Dickey-Fuller test statistic is used :

$$\Delta \hat{u}_t = \phi \hat{u}_{t-1} + \sum_{i=1}^{p} \delta_i \Delta \hat{u}_{t-i} + v_t \tag{1.10}$$

The ADF-statistic is obtained then as the DF-t-statistic for ϕ in the regression of (1.10).

The DF-t-statistics for ϕ are given in table 1.4. Note that only results for p=0 are reported, so that we can speak about an ordinary DF-test. This can be justified by the fact that increasing the order of p did not seem to have influence on the test outcomes. The critical values for the DF-test statistic, with T=100 and the number of variabels, K=2, are equal to 3,37 if α=5% and 4,07 if α=1%. The critical values for T=200 and K=2 are 3,37 if α=5% and 4,00 if α=1% (see Engle and Yoo,1987). We notice that the critical values are greater than the conventional critical values 1,96 if α=5% and 2.58 if α=1%. The reason is that the conventional critical values lead too often to a rejection of the nulhypothesis. Because no critical values for T=500 are tabulated (T=200 is the highest number for T for which critical values are calculated), we take, for simplicity, that the critical value is 3,5, so that the linear combination is stationary if $|t|>3,5$ and the linear combination is non-stationary if $|t|<3,5$.
From the results of table 1.4, we can conclude that no co-integrating relation exists between the exchange rate series and relative consumer price indices, both for weekly and monthly data. This means that PPP does not hold in the long-run; there is no equilibrium relationship found between the exchange rate series and relative consumer price indices.

2. The Uncovered Interest Parity Hypothesis

In this section, the Uncovered Interest Parity(UIP) as a long-run hypothesis will be tested. For this purpose the Engle and Granger methodology (Engle and Granger, 1987), which we have already applied to the PPP in the previous section, will be used. The Uncovered Interest Parity Hypothesis states that a nominal interest rate differential of bonds, denominated in different currencies, equals the expected change in the exchange rate. Moreover, for the validity of UIP, the following assumptions have to be fulfilled : capital has to be perfectly mobile and domestic and foreign bonds are perfect substitutes, which implies that there are no transaction costs, no differences in national tax systems on capital income, and no risk premia in forward markets, which are in addition regarded as efficient.

Formally, the UIP can be expressed by the following equation :

$$r_t - r_t^* = E_t(s_{t+1}) - s_t \qquad (2.1)$$

with :

r_t : domestic nominal interest rate

r_t^* : foreign nominal interest rate

s_t : the (logarithm of the) exchange rate (foreign currency expressed in domestic currency)

$E_t(.)$: conditional expectation, given all information available up to time period t

Concerning the data, weekly data (wednesday data) for the euro-rate on three months for Belgium, the Netherlands, Germany and the USA will be used [1]. The interest rate series are shown in figures 2.1 to 2.8.

There exist different methods to test the validity of UIP. Kirchgässner and Wolters (1989) tested the validity of UIP by using the methodology proposed by Engle and Granger (1987). Here, we will follow the same procedure. Kirchgässner and Wolters start from the asset market approach to explain movements of the exchange rate, being the relative price of two durable assets in an efficient market. This asset market approach shows that, under rational expectations, (short-run) changes of exchange rates are uncorrelated, so that the expectations of these changes are assumed to follow a stationary process. Consequently, equation (2.1) can be rewritten as :

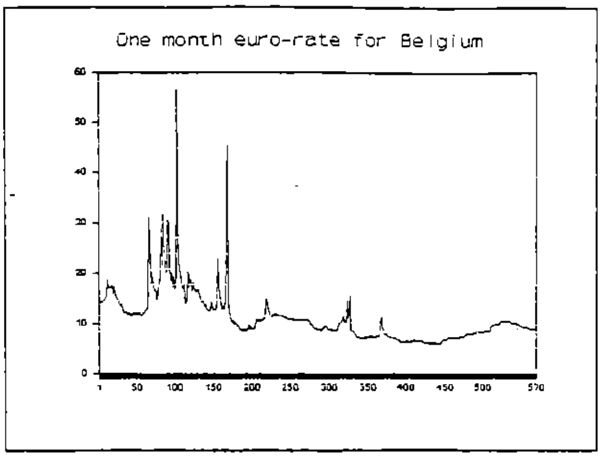
$$r_t - r_t^* = \eta_t \tag{2.2}$$

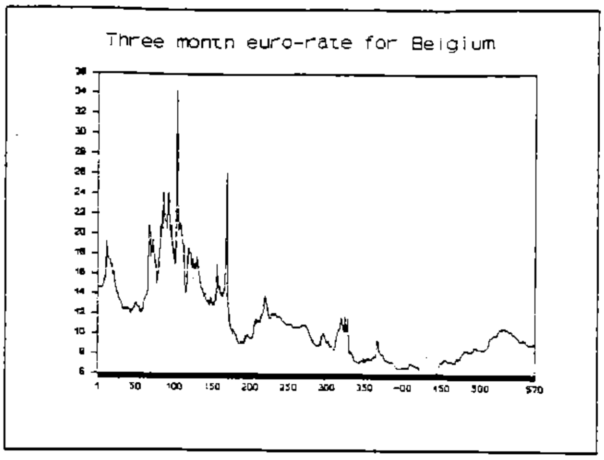
with η_t stationary.

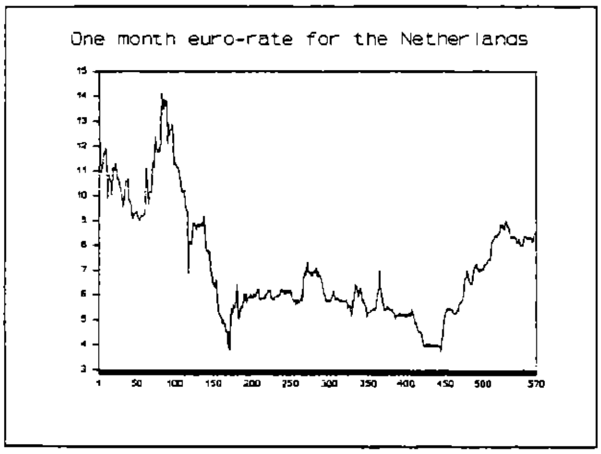
Thus, we have to test whether the interest rate differential is stationary. When the interest rates are stationary, there wil be no problem. However, when we perform tests (see equations (1.3) to

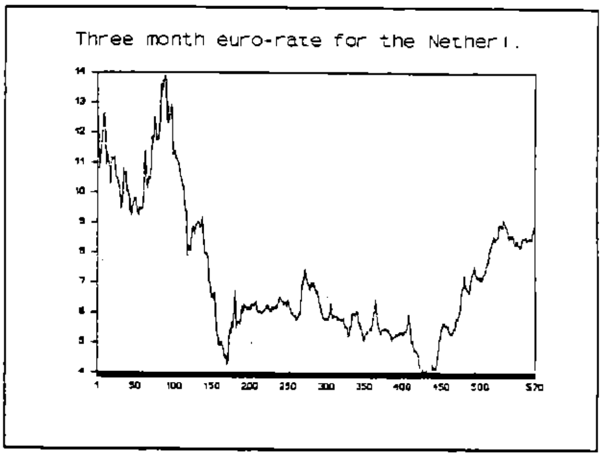
[1] Note that the analysis was also performed for the euro-rate on one month for the concerning countries. Nevertheless, we will only report here the outcomes of the stationarity and cointegrating tests for the euro-rate on three months, because the results of the tests for the euro-rate on one month and three months were very similar.

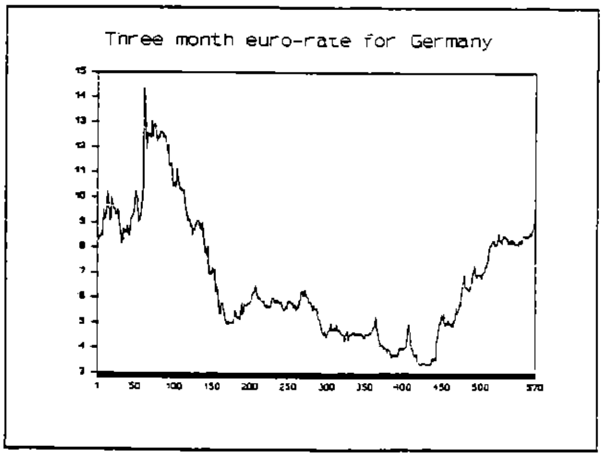
Figures 2.1 - 2.8

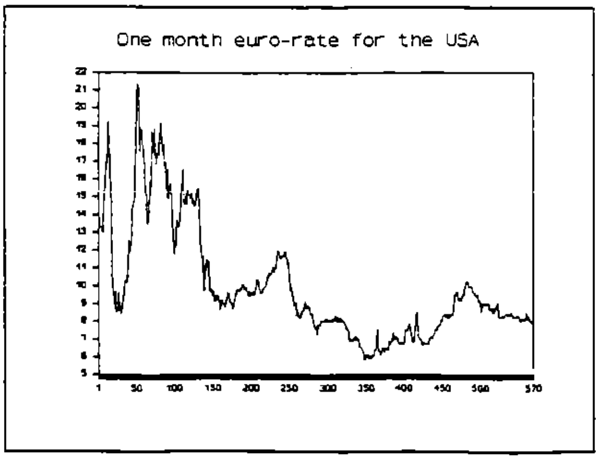


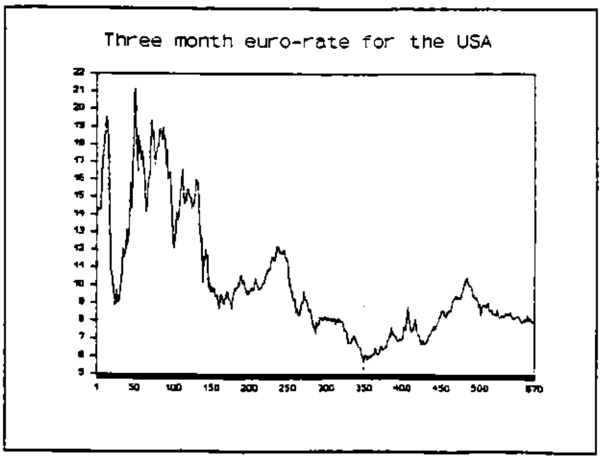












(1.8)) for stationarity on the interest rate, the outcomes (see table 2.1 of the appendix [2]) indicates that all the series for interest rates are integrated of order one, i.e., the series become stationary after differencing once. As a consequence, we have to test whether the interest rate differentials are stationary. The results are tabulated in table 2.2. The results of table 2.2 indicate that for weekly data all interest rate differentials are stationary with significance level α=0,01% with exception of the interest rate differentials between the Netherlands on one side and the USA on the other side. For these last differentials it depends on which test and which significance level is chosen whether the interest rate differential is stationary .
For monthly data, it depends on which test and which significance level is chosen whether the interest rate differential is stationary.

In a following step, it will be investigated whether the interest rates are cointegrated, i.e., there exists a long-run relationship between them. First, we estimate the cointegrating relation :

$$r_t = \alpha + \beta r_t^* + u_t \tag{2.3}$$

with r_t and r_t^* integrated of order one.

Secondly, we test with a DF-test statistic whether the estimated residuals $\hat{u}_t$ are stationary (see eq.(1.10)). The results are

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[2] Note that we only reported the results for the alternative test because the results of both tests were similar.

tabulated in table 2.3 of the appendix. The results show that there exists a cointegrating relation between all interest rates with exception between $r3_{nlg}$ and $r3_{usd}$ for weekly data and between $r3_{bef}$ and $r3_{usd}$, between $r3_{nlg}$ and $r3_{usd}$ and between $r3_{bef}$ and $r3_{dem}$ for monthly data. These results are rather surprising because we expected more coherence between the exchange rates and their corresponding interest rate differentials for monthly data.

3. Testing the direction of causality and developing a transfer function for the estimation of an exchange rate equation

In a first paragraph, the direction of causality between the exchange rates on the one side and the relative consumer price indices and the interest rate differentials on the other hand is tested. On the basis of the outcomes of the causality tests a transfer function to estimate an exchange rate equation will be set up in a second paragraph.

3.1. Testing the direction of causality

The aim of this section is to find out whether causality runs from the relative consumer price indices and the interest rate differentials to the exchange rate or the other way around. The concept of causality was first elaborated by Granger(1969). Granger's theory of causality is entirely based on the assumption that the future cannot cause the past. By definition (see Granger,1969), it is said that Y_t is causing X_t, denoted by $Y_t \rightarrow X_t$ and with Y_t and X_t stationary time series, if the predictions of X_t are better using all available information than if the information apart from Y_t had been used. If X_t is causing Y_t and also Y_t is causing X_t, feedback is said to occur, denoted by $Y_t \Leftrightarrow X_t$. To test causality between the exchange rates on the one side and the relative consumer price indices and the interest differentials on the other hand, we use Geweke's test, which is based on Sims'test.

Sims (1972) showed that unidirectional causality can be tested by regressing Y_t on past and future values of X_t, given by :

$$Y_t = \sum_{i=n}^{-m} \gamma_{-i} X_{t-i} + v_t \tag{3.1}$$

If causality runs from X to Y only, future values of X in the regression should have coefficients insignificantly different from zero as a group, which can be tested by an F-test with as nullhypothesis that Y does <u>not</u> cause X (or X only causes Y), i.e., $H_0 : \gamma_1 = \gamma_2 = \ldots = \gamma_m = 0$.

Generally v_t is serially correlated, so that a feasible generalized least squares estimator or a prefilter have to be employed to cope with the serial correlation in v_t. A variant of the Sims' test, developed by Geweke (1983), avoids such a feasible generalized least squares estimator or prefilter, assuming that v_t can be approximated by an AR(p)-proces. The Sims test can then be rewritten as :

$$Y_t = \sum_{j=1}^{p} \phi_{-j} Y_{t-j} + \sum_{i=n+p}^{-m} \gamma_{-i} X_{t-i} \tag{3.2}$$

As above, the hypothesis that Y does not cause X can be tested then by an F-statistic with the nullhypothesis $H_0 : \gamma_1 = \gamma_2 = \ldots = \gamma_m = 0$.

The outcomes of Geweke's test to test the direction of causality between the exchange rates on the one side and the relative consumer price indices and the interest differentials on the other hand are given in tables 3.1 and 3.2.

The results of table 3.1 show that for weekly data causality is running from the relative consumer price indices to the exchange rate returns with exception of the Belgian Franc vis à vis the DM. For the Dutch Guilder vis à vis the DM even simultaneous feedback seems to exist. For monthly data only causality is running from the relative consumer price index to the exchange rate for the Belgian franc vis à vis the DM. Between the interest rate differentials and the exchange rate returns (see table 3.2) causality is running from the Belgian Franc vis à vis the USD to the corresponding interest rate differential for weekly data and from the interest rate differential between the Netherlands and the USA to the corresponding exchange rate for monthly data. Moreover, simultaneous feedback exists between the Belgian Franc vis à vis the DM and their corresponding interest rate differential for weekly data only and between the Dutch Guilder vis à vis the DM and their corresponding interest rate differentials for both weekly and monthly data. This implies a room for simultaneous explanation of exchange rate returns and interest rates. Such a simultaneous explanation will be explored in a following paper.

3.2. Development of a transfer function for the estimation of an exchange rate equation

In this paragraph transfer functions for the corresponding exchange rates are constructed with as input variables the relative consumer price indices and the interest rates differentials. The construction and use of a transfer function consists of 3 steps :

1. The identification phase : an ARIMA-model for the input variables is estimated, and the same ARIMA-model is applied to the output variables (prewhitening). The crosscorelations between the prewhitened input and output variables are calculated then to get insight into the structure of the transfer function.
2. The estimation phase : the coefficients of the specified models are estimated (by conditional least squares or maximum likelihood) and the adequacy of the model is analysed.
3. The forecasting phase : the estimated models are used to generate forecasts.

In the next alineas, firstly, the univariate ARIMA-models for each exchange rate are reported, followed by the estimation results of the transfer functions. Next, the absolute and relative Theil's inequality indices are calculated in order to controle if the forecasted values approximate well the observed values. The absolute and relative indices are given by :

$$T_1 := \frac{\sqrt{\sum_{t=1}^{T} (y_{it} - y_{it})^2}}{\sqrt{\sum_{t=1}^{T} y_{it}^2}} \qquad T_2 = \frac{\sqrt{\sum_{t \in T_f} (y_{it} - y_{it})^2}}{\sqrt{\sum_{t=1} y_{it}^2} + \sqrt{\sum_{t=1}^{T} y_{it}^2}} \quad (3.3)$$

respectively with y_{it} the observed value at period t, $\hat{y}_{it}$ the relating forecasted value and T_f the forecasting period. Theil's inequality index is preferably below 0.4 for good forecasts, while the absolute index is always between zero and one.

Let us now first look at the univariate ARIMA-models of each exchange rate, which are given by [1] :

* BF/USD
  - weekly data

$$\underset{(-1,89)}{(1 + 0{,}080L^{6})}\underset{(1{,}72)}{(1 - 0{,}073L^{32})}\Delta\hat{s}_t = \underset{(-2{,}68)}{(1 + 0{,}113L^{2})}\underset{(-2{,}49)}{(1 + 0{,}106L^{41})}\,x$$

$$\underset{(-3{,}90)}{(1 + 0{,}166L^{48})}\hat{\epsilon}_t \qquad (3.4)$$

with residual standard error : 0.01501563

  - monthly data

$$\underset{(3{,}61)}{(1 - 0{,}302L} \underset{(2{,}09)}{- 0{,}175L^{3})}\Delta\hat{s}_t = \underset{(-2{,}13)}{(1 + 0{,}194L^{14})}\hat{\epsilon}_t \qquad (3.5)$$

with residual standard error : 0,02789168

* NLG/USD
  - weekly data

---

[1] Estimated t-values (ratio between estimated parameters and standard errors) are given between brackets.

$$(1 - \underset{(2,72)}{0,114}L^2 + \underset{(-2,29)}{0,096}L^6)(1 - \underset{(2,64)}{0,112}L^{41} - \underset{(2,84)}{0,121}L^{48})\Delta\hat{s}_t =$$

$$(1 + \underset{(-2,15)}{0,092}L^8)(1 + \underset{(-1,86)}{0,079}L^{23})\hat{\epsilon}_t \qquad (3.6)$$

with residual standard error : 0.01500273

- <u>monthly data</u>

$$(1 - \underset{(3,83)}{0,323}L)\Delta\hat{s}_t = (1 + \underset{(-1,98)}{0,182}L^{14})\hat{\epsilon}_t \qquad (3.7)$$

with residual standard error : 0,02785555

* <u>BF/DM</u>
  - <u>weekly data</u>

$$(1 - \underset{(2,29)}{0,097}L^9)(1 - \underset{(1,96)}{0,084}L^{50})L^{50}\Delta\hat{s}_t = (1 - \underset{(4,69)}{0,198}L^{10})(1 + \underset{(-3,15)}{0,133}L^{12})x$$

$$(1 + \underset{(-4,69)}{0,193}L^{16} + \underset{(-4,11)}{0,172}L^{20})(1 - \underset{(2,37)}{0,101}L^{47})\hat{\epsilon}_t \qquad (3.8)$$

with residual standard error : 0,00374869

- <u>monthly data</u>

with residual standard error : 0,00678325

$$(1 - \underset{(4,34)}{0,350}L - \underset{(2,89)}{0,235}L^{12})\Delta\hat{s}_t = (1 + \underset{(-2,57)}{0,224}L^{5})\hat{\varepsilon}_t \qquad (3.9)$$

* <u>NLG/DM</u>

- <u>weekly data</u>

$$(1 + \underset{(-2,03)}{0,086}L^{16})(1 + \underset{(-2,49)}{0,105}L^{32} - \underset{(2,46)}{0,103}L^{45})\Delta\hat{s}_t =$$

$$(1 - \underset{(2,72)}{0,113}L^{23} - \underset{(2,34)}{0,098}L^{24} + \underset{(-2,34)}{0,098}L^{43})\hat{\varepsilon}_t \qquad (3.10)$$

with residual standard error : 0,00164963

- <u>monthly data</u>

$$(1 - \underset{(4,01)}{0,365}L)(1 + \underset{(-2,69)}{0,253}L^{2})(1 + \underset{(-2,18)}{0,197}L^{6})\Delta\hat{s}_t = (1 - \underset{(2,36)}{0,215}L^{5})\,x$$

$$(1 + \underset{(-2,42)}{0,224}L^{10})(1 + \underset{(-2,52)}{0,228}L^{21})(1 + \underset{(-2,52)}{0,227}L^{24})\hat{\varepsilon}_t \qquad (3.11)$$

with residual standard error : 0,00245776

After identification, with the help of the crosscorrelations, the following transferfunctions were constructed and estimated :

* <u>BF/USD</u>

- <u>weekly data</u>

$$\underset{(-2{,}44)}{(1 + 0{,}108L^{6})}\underset{(2{,}12)}{(1 - 0{,}095L^{32})}\Delta\hat{s}_t = (\underset{(2{,}66)}{4{,}122L^{30}} + \underset{(-2{,}31)}{3{,}614L^{38}} +$$

$$\underset{(-1{,}63)}{2{,}551L^{43}})\Delta^2(p_t - p_t^*) + (\underset{(-4{,}66)}{-0{,}0023} + \underset{(-3{,}46)}{0{,}0017L^{1}})(r_t - r_t^*) +$$

$$\underset{(-1{,}79)}{(1 + 0{,}079L^{2})}\underset{(-2{,}73)}{(1 + 0{,}123L^{41})}\underset{(-3{,}95)}{(1 + 0{,}177L^{48})}\hat{\epsilon}_t \qquad (3.12)$$

with residual standard error : 0,01471098

- <u>monthly data</u>

$$(1 + \underset{(3{,}95)}{0{,}331L} - \underset{(2{,}48)}{0{,}206L^{3}})\Delta\hat{s}_t = \underset{(-4{,}03)}{-0{,}009}\Delta(r_t - r_t^*) +$$

$$\underset{(-1{,}87)}{(1 + 0{,}173L^{14})}\hat{\epsilon}_t \qquad (3.13)$$

with residual standard error : 0,02633769

* <u>NLG/USD</u>
  - <u>weekly data</u>

$$(1 - \underset{(2{,}32)}{0{,}104L^{2}} + \underset{(-2{,}04)}{0{,}091L^{6}})(1 - \underset{(2{,}17)}{0{,}097L^{41}} - \underset{(3{,}30)}{0{,}148L^{48}})\Delta\hat{s}_t =$$

$$(\underset{(1{,}61)}{2{,}076L^{38}} + \underset{(-2{,}03)}{2{,}664L^{42}} + \underset{(-1{,}59)}{2{,}106L^{47}})\Delta^2(p_t - p_t^*) +$$

$$(-0{,}0078L - 0{,}0033L^{18} - 0{,}0022L^{44})\Delta(r_t - r_t^*) +$$
$$\quad (-4{,}71) \qquad (2{,}19) \qquad (1{,}74)$$

$$(1 + 0{,}086L^{8})(1 + 0{,}076L^{23})\hat{\epsilon}_t \qquad (3.14)$$
$$\quad (-1{,}91) \qquad (-1{,}69)$$

with residual standard error : 0,01476698

- monthly data

$$(1 - 0{,}334L)\Delta\hat{s}_t = (-0{,}0077L^{2} - 0{,}0069L^{14} -$$
$$\quad (3{,}68) \qquad (-1{,}86) \qquad (2{,}66)$$

$$0{,}0059L^{16})\Delta(r_t - r_t^*) + \hat{\epsilon}_t \qquad (3.15)$$
$$(2{,}48)$$

with residual standard error : 0,02659652

* BF/DM

  - weekly data

$$(1 - 0{,}138L^{3})(1 + 0{,}071L^{17} - 0{,}070L^{20} - 0{,}110L^{45})(1 - 0{,}152L^{50})\Delta\hat{s}_t =$$
$$\quad (3{,}07) \qquad (-1{,}56) \qquad (1{,}54) \qquad (2{,}43) \qquad (3{,}35)$$

$$(1{,}005L^{24} + 2{,}244L^{33} + 1{,}39L^{37})\Delta^{2}(p_t - p_t^*) +$$
$$\quad (2{,}79) \qquad (-5{,}90) \qquad (-3{,}72)$$

$$(-0{,}0012 + 0{,}0007L - 0{,}005L^{9} + 0{,}0018L^{10} - 0{,}0008L^{11} +$$
$$(-10{,}77) \qquad (-6{,}55) \qquad (4{,}63) \qquad (-14{,}05) \qquad (7{,}27)$$

$$0{,}0002L^{21} - 0{,}0001L^{42})(r_t - r_t^*) + (1 + 0{,}143L^4)(1 + 0{,}105L^{12})$$
$$\quad(-2{,}44) \qquad (2{,}12) \qquad\qquad\qquad (-3{,}16) \qquad (-2{,}32)$$

$$(1 - 0{,}084L^{18} - 0{,}087L^{37} - 0{,}116L^{49})\hat{\epsilon}_t \qquad (3.16)$$
$$\quad(1{,}87) \qquad (1{,}92) \qquad (2{,}56)$$

with residual standard error : 0,00302099

- monthly data

$$(1 - 0{,}560L)\Delta\hat{s}_t = (-0{,}409L^{10} - 0{,}434L^{11})(p_t - p_t^*) +$$
$$\quad(6{,}69) \qquad\qquad (-2{,}55) \qquad (2{,}77)$$

$$(0{,}0059L^{11} + 0{,}002L^{24})\Delta(r_t - r_t^*) + (1 + 0{,}263L^{15})\hat{\epsilon}_t \qquad (3.17)$$
$$\quad(7{,}95) \qquad (-2{,}72) \qquad\qquad (-2{,}54)$$

with residual standard error : 0,00571292

* NLG/DM

- weekly data

$$(1 + 0{,}099L^{29})(1 - 0{,}155L^{45})\Delta\hat{s}_t = 0{,}148\Delta^2(p_t - p_t^*) +$$
$$\quad(-2{,}23) \qquad (3{,}51) \qquad\qquad (1{,}11)$$

$$0{,}0006L^{26} + 0{,}0006L^{28} + 0{,}0009L^{32})\Delta(r_t - r_t^*) +$$
$$\quad(2{,}55) \qquad (-2{,}33) \qquad (-3{,}66)$$

$$(- \underset{(-7,55)}{0,0019} - \underset{(3,28)}{0,0008L} - \underset{(3,77)}{0,001L^2} - \underset{(2,85)}{0,0007L^{14}} - \underset{(2,33)}{0,0006L^{15}} -$$

$$(1 - \underset{(1,71)}{0,075L^{23}} - \underset{(2,25)}{0,099L^{24}})(1 - \underset{(2,53)}{0,112L^{42}})\hat{\varepsilon}_t \qquad (3.18)$$

with residual standard error : 0,00149463

- <u>monthly data</u>

$$(1 - \underset{(2,13)}{0,225L})(1 + \underset{(-3,01)}{0,328L^2})\Delta\hat{s}_t = (\underset{(-3,93)}{-0,0031L} - \underset{(3,59)}{0,0018L^{12}} +$$

$$\underset{(-4,25)}{0,0018L^{19}} + \underset{(-3,21)}{0,0015L^{20}} + \underset{(-3,47)}{0,0014L^{23}} - \underset{(2,28)}{0,001L^{24}})\Delta(r_t - r_t^*)$$

$$(1 + \underset{(-3,98)}{0,400L^{10}})(1 + \underset{(-1,86)}{0,196L^{12}})\hat{\varepsilon}_t \qquad (3.19)$$

with residual standard error : 0,0015973

We remark that for the Belgian Franc and the Dutch Guilder vis à vis the USD and the Dutch Guilder vis à vis the DM the relative consumer price indices were not significant in the transferfunctions for monthly data. This fits with the outcomes of the causality tests from the relative consumer price indices to their corresponding exchange rates, namely that only causality was running from the relative consumer price index between Belgium and Germany to the Belgian Franc vis à vis the DM.

In table 3.3.a and 3.3.b Theil's inequality indices for forecasting periods of 8 and 22 weeks and 5 months respectively are tabulated. The results of table 3.3.a (weekly figures) firstly shows the natural result that the shorter the forecasting period is, the better the forecasts are. Secondly, the transfer function generates the best forecasts for the Belgian Franc vis à vis the DM, while for the Belgian Franc vis à vis the US Dollar and for the Dutch Guilder vis à vis the DM the random walk model and the ARIMA-model respectively are superior in forecasting. Nevertheless, for the Dutch Guilder vis à vis the DM, the calculated Theil's inequality indices are very low, so that in fact every model can be accepted. For the Dutch Guilder vis à vis the US Dollar the transfer function generate the best forecasts for a sampling period of 8 weeks, while the random walk model is better for a sampling period of 22 weeks. Finally, we notice that the random walk model always is superior to the ARIMA-model in forecasting with exception of the Dutch Guilder vis à vis the DM. Table 3.3.b (monthly figures) indicates that the Belgian franc and the Dutch guilder vis à vis the US Dollar follow a random walk pattern, while the Belgian franc and the Dutch Guilder vis à vis the DM are predicted best by means of an ARIMA-model. Exchange rate returns [2] are not predicted well for every model. For example, for the Belgian Franc vis à vis the DM, we have for the absolute Theil's inequality index, T_1, for a forecasting period of 8 and 22 weeks 0,882719 and 0,910414 for the ARIMA-model and 0,801056 and 1,038265 for the transfer function respectively. For a forecasting period of 9 months the corresponding figures are 0,924637 for the ARIMA-model and 0,757810 for the transfer function.

[2] Note that for exchange rate returns Theil's inequality index is equal to 1 in the case of random walk.

4. Conclusion

Starting from the results of statistical test on the stationarity and the long-run validity of PPP and UIP for exchange rates, an ARIMA-model and a transfer function model for each exchange rate (return) was constructed, estimated and used to generate forecasts. By means of the absolute and relative Theil's inequality indices, the forecasted values were evaluated and compared to the random walk model. For weekly data, only for one exchange rate (return), i.e. the Belgian Franc vis à vis the DM, the transfer function model, with price and interest rate differentials, was superior to the ARIMA-model and random walk model in forecasting. For the Belgian Franc vis à vis the US Dollar and the Dutch Guilder vis à vis the DM, the random walk and the ARIMA-model respectively are superior w.r.t. the PPP and UIP-hypothesis. For monthly data, the random walk model predicted best for the Belgian Franc and the Dutch Guilder vis à vis the US Dollar, while the ARIMA-model was superior for the Belgian Franc and the Dutch Guilder vis à vis the DM. Forecasting performance for exchange rate models differencing from the PPP and UIP-hypothesis and simultaneous exploration of exchange rate (returns) and interest rates will be examined in a following paper.

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APPENDIX

Table 1.1 : Results of the F-test for the strict PPP-hypothesis

Currency	dep. var.	const.	$p_t-p^*_t$	$\Delta(p_t-p^*_t)$	R^2	DW	F-test
F-test with weekly data : T=570							
BEF/USD	s_t	3,629 (421,000)	4,306 (21,048)	-	0,438	0,009	286004,031
	Δs_t	0,000 (0,217)	-	-0,207 (-0,286)	0,000	1,861	2,929
NLG/USD	s_t	1,018 (86,954)	1,628 (15,296)	-	0,292	0,008	18214,826
	Δs_t	0,000 (-0,659)	-	-0,558 (-0,807)	0,001	1,897	5,220
BEF/DM	s_t	2,826 (908,730)	1,238 (54,076)	-	0,837	0,012	3306492,000
	Δs_t	0,000 (2,066)	-	0,114 (0,509)	0,000	1,951	15,968
NLG/DM	s_t	0,112 (197,690)	0,053 (1,111)	-	0,002	0,019	44580,645
	Δs_t	0,000 (0,574)	-	0,029 (0,383)	0,000	1,865	166,766
F-test with monthly data : T=131							
BEF/USD	s_t	3,95 (159,89)	4,392 (10,557)	-	0,463	0,045	71646,859
	Δs_t	0,0008 (0,314)	-	0,357 (0,540)	0,002	1,313	1,127
NLG/USD	s_t	0,920 (57,754)	1,632 (7,493)	-	0,303	0,034	3737,959
	Δs_t	-0,002 (-0,643)	-	-0,450 (-0,734)	0,004	1,311	5,771
BEF/DM	s_t	3,012 (858,54)	1,226 (26,989)	-	0,850	0,055	906878,063
	Δs_t	0,002 (2,374)	-	0,085 (0,448)	0,002	1,301	23,472
NLG/DM	s_t	0,114 (79,194)	0,147 (1,67)	-	0,021	0,059	14032,664
	Δs_t	0,0002 (0,652)	-	0,047 (0,785)	0,005	1,299	255,203

(*) The critical values of the F-test, i.e. F(2,500) for weekly data and F(2,120) for monthly data, are equal to 3,014 and 3,072 respectively if α=0,05 and are equal to 4,713 and 4,787 respectively if α=0,01.

Table 1.2 : Stationarity tests for the exchange rates and relative consumer price indices : variables are in first order differences of logarithms

DF-test statistic : $\hat{\tau}$

Currency		$\hat{\tau}_\epsilon$	$\hat{\tau}_\mu$		$\hat{\tau}_\tau$	
DF-test statistic with weekly data : T=568						
BEF/USD	Δs_t	-22,216	-22,199	-	-22,485	c t
	$\Delta(p_t - p^*_t)$	-5,387	-5,422	-	-5,419	- -
NLG/USD	Δs_t	-22,598	-22,583	-	-84,260	- t
	$\Delta(p_t - p^*_t)$	-6,044	-6,670	c	-6,669	c -
BEF/DEM	Δs_t	-22,973	-23,200	c	-23,226	c -
	$\Delta(p_t - p^*_t)$	-5,495	-5,951	c	-5,940	c -
NLG/DEM	Δs_t	-22,292	-22,282	-	-22,293	- -
	$\Delta(p_t - p^*_t)$	-6,249	-6,243	-	-6,226	- -
DF-test statistic with monthly data : T=129						
BEF/USD	Δs_t	-7,849	-7,819	-	-8,455	c t
	$\Delta(p_t - p^*_t)$	-7,683	-7,705	-	-7,872	- -
NLG/USD	Δs_t	-7,919	-7,898	-	-8,325	- t
	$\Delta(p_t - p^*_t)$	-8,264	-9,540	c	-9,617	- -
BEF/DEM	Δs_t	-7,447	-7,751	-	-8,069	c -
	$\Delta(p_t - p^*_t)$	-7,474	-8,545	c	-8,797	c -
NLG/DEM	Δs_t	-7,935	-7,928	-	-7,902	- -
	$\Delta(p_t - p^*_t)$	-9,039	-9,004	-	-9,065	- -

(*) A "c" or a "t" in the table indicates that the constant term and the coefficient of the deterministic time trend respectively are significant with confidence level $1-\alpha=0,95$.

Table 1.3 : Critical values for the DF-statistic for T=500

	Sample Size	Empirical cumulative distribution of $\hat{\tau}$ for $\rho=1$ Probability of a smaller value			
	T	0,01	0,025	0,05	0,10
$\hat{\tau}_{\epsilon}$	100	-2,60	-2,24	-1,95	-1,61
	500	-2,58	-2,23	-1,95	-1,62
$\hat{\tau}_{\mu}$	100	-3,51	-3,17	-2,89	-2,58
	500	-3,44	-3,13	-2,87	-2,57
$\hat{\tau}_{\tau}$	100	-4,04	-3,73	-3,45	-3,15
	500	-3,98	-3,68	-3,42	-3,13

(*) W.A. FULLER(1976), Introduction to Statistical Time Series, New York, Wiley, p.373

Table 1.4 : The DF-test statistic

Currency	DF-statistic	
	weekly data	monthly data
BEF/USD	-2,570	-2,656
NLG/USD	-2,411	-2,360
BEF/DEM	-1,231	-1,324
NLG/DEM	-1,783	-1,385

Table 2.1 : Stationarity tests for the interest rates : variables are in differences of %

DF-test statistic : $\hat{\tau}$

	$\hat{\tau}_{\epsilon}$	$\hat{\tau}_{\mu}$	$\hat{\tau}_{\tau}$
DF-test statistic with weekly data : T=568			
R3BEF	-30,081	-30,057 -	-30,031 - -
R3NLG	-21,957	-21,952 -	-21,978 - -
R3DEM	-23,108	-23,089 -	-23,095 - -
R3USD	-20,787	-20,777 -	-20,760 - -
DF-test statistic with monthly data : T=129			
R3BEF	-8,424	-8,425 -	-8,423 - -
R3NLG	-8,090	-8,079 -	-8,327 - -
R3DEM	-8,554	-8,520 -	-8,630 - -
R3USD	-8,562	-8,547 -	-8,525 - -

(*) A "c" or a "t" in the table indicates that the constant term and the coefficient of the deterministic time trend respectively are significant with confidence level 1-α=0,95.

Table 2.2 : Stationarity tests for the interest rate differentials : $r_t - r^*_t$

DF-test statistic : $\hat{\tau}$

r_t	r^*_t	$\hat{\tau}_\epsilon$	$\hat{\tau}_\mu$		$\hat{\tau}_\tau$	
DF-test statistic with weekly data : T=569						
R3BEF	R3USD	-5,968	-6,378	c	-6,396	c -
R3NLG	R3USD	-1,662	-2,446	-	-3,085	c -
R3BEF	R3DEM	-3,015	-5,630	c	-6,931	c t
R3NLG	R3DEM	-4,185	-4,717	c	-4,729	- -
DF-test statistic with monthly data : T=130						
R3BEF	R3USD	-3,335	-3,374	-	-3,369	- -
R3NLG	R3USD	-1,777	-2,641	-	-3,753	c t
R3BEF	R3DEM	-1,527	-1,882	-	-3,020	c t
R3NLG	R3DEM	-3,524	-3,760	-	-3,775	- -

(*) A "c" or a "t" in the table indicates that the constant term and the coefficient of the deterministic time trend respectively are significant with confidence level $1-\alpha=0,95$.

Table 2.3 : Estimation of equation (2.3) and the DF-test statistic

r_t	r^*_t	$\hat{\alpha}$	$\hat{\beta}$	R^2	DW	DF-statistic
Weekly data						
R3BEF	R3USD	1,517 (4,633)	0,939 (30,731)	0,624	0,265	-6,353
R3NLG	R3USD	1,998 (10,365)	0,521 (28,963)	0,596	0,042	-2,465
R3BEF	R3DEM	2,392 (7,694)	1,286 (29,659)	0,608	0,242	-5,890
R3NLG	R3DEM	1,328 (14,347)	0,885 (68,546)	0,892	0,109	-5,460
Monthly data						
R3BEF	R3USD	2,909 (6,619)	0,743 (18,199)	0,720	0,272	-3,101
R3NLG	R3USD	1,990 (5,089)	0,520 (14,317)	0,614	0,198	-2,484
R3BEF	R3DEM	3,590 (8,240)	1,024 (16,840)	0,687	0,141	-1,890
R3NLG	R3DEM	1,303 (7,056)	0,889 (34,488)	0,902	0,207	-4,257

Table 3.1 : Geweke's F-test for the direction of causality between the relative consumption price index and the exchange rate

Currency	$Y_t=\Delta(p_t-p^*_t)$ $X_t=\Delta s_t$	$Y_t=\Delta s_t$ $X_t=\Delta(p_t-p^*_t)$
Geweke's F-test with weekly data		
BEF/USD	4,883	1,292
NLG/USD	4,480	0,074
BEF/DEM	0,142	0,095
NLG/DEM	3,720	3,333
Geweke's F-test with monthly data		
BEF/USD	1,067	1,581
NLG/USD	0,664	0,043
BEF/DEM	5,179	1,196
NLG/DEM	1,301	1,622

(*) The values of p, n and m were set equal to 1, 1 and 2 respectively.

(**) The critical F-values, F(2,500) for weekly data and F(2,120) for monthly data, with confidence level $1-\alpha=0,95$ are equal to 3,014 and 3,072 respectively.

Table 3.2 : Geweke's F-test for the direction of causality between the interest rate differential and the exchange rate

r_t	r^*_t	$Y_t=\Delta(r_t-r^*_t)$ $X_t=\Delta s_t$	$Y_t=\Delta s_t$ $X_t=\Delta(r_t-r^*_t)$
Geweke's F-test with weekly data			
R3BEF	R3USD	0,470	4,791
R3NLG	R3USD	2,832	1,861
R3BEF	R3DEM	3,135	5,161
R3NLG	R3DEM	8,739	16,264
Geweke's F-test with monthly data			
R3BEF	R3USD	2,502	0,064
R3NLG	R3USD	4,117	1,080
R3BEF	R3DEM	0,053	15,379
R3NLG	R3DEM	34,362	10,361

(*) The values of p,n and m were set equal to 1, 1 and 2 respectively.

(**) The critical F-values, F(2,500) for weekly data and F(2,120) for monthly data, with confidence level $1-\alpha=0,95$ are equal to 3,014 and 3,072 respectively.

Table 3.3.a : Theil's inequality coefficient with a forecasting period of 8 and 22 weeks respectively and the residual sampling standard errors

Currency		Arima	Transfer	Random Walk
Forecasting period : 8 weeks				
BEF/USD	T_1	0,024748	0,024060	0,022883
	T_2	0,012500	0,012151	0,011545
NLG/USD	T_1	0,026449	0,023342	0,023517
	T_2	0,013365	0,011783	0,011864
BEF/DEM	T_1	0,002453	0,001374	0,002059
	T_2	0,001226	0,000687	0,001029
NLG/DEM	T_1	0,000304	0,000394	0,000339
	T_2	0,000152	0,000197	0,000169
Forecasting period : 22 weeks				
BEF/USD	T_1	0,086680	0,087655	0,078834
	T_2	0,044720	0,045229	0,040534
NLG/USD	T_1	0,086953	0,084820	0,080568
	T_2	0,044891	0,043706	0,041460
BEF/DEM	T_1	0,004809	0,001266	0,002998
	T_2	0,002399	0,000633	0,001497
NLG/DEM	T_1	0,000309	0,000673	0,000630
	T_2	0,000154	0,000336	0,000314
Residual standard errors for exchange rates				
BEF/USD		0,01563015	0,01471098	
NLG/USD		0,01500273	0,01476698	
BEF/DEM		0,00374869	0,00302099	
NLG/DEM		0,00164963	0,00149463	

Table 3.3.b : Theil's inequality coefficient with a forecasting period of 5 months and the residual standard errors

Currency		Arima	Transfer	Random Walk
Forecasting period : 5 months				
BEF/USD	T_1	0.066712	0.066055	0.030332
	T_2	0.034290	0.033931	0.015133
NLG/USD	T_1	0.059124	0.058242	0.030966
	T_2	0.030289	0.029824	0.015435
BEF/DEM	T_1	0.001998	0.003359	0.003267
	T_2	0.000998	0.001677	0.001635
NLG/DEM	T_1	0.000624	0.000767	0.000888
	T_2	0.000312	0.000383	0.000444
Residual standard errors for exchange rates				
BEF/USD		0.02789168	0.02633769	
NLG/USD		0.02785555	0.02659652	
BEF/DEM		0.00678325	0.00571292	
NLG/DEM		0.00245776	0.0015973	

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