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Non-Stationary Interest Rate Differentials and the Role of Monetary Policy

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Abstract

The present work deals with a frequently detected failure of the uncovered interest rate parity (UIP) - the absence of bivariate cointegration between domestic and foreign interest rates. We explain non-stationarity of the interest differential via central bank reactions to exchange rate variations. Thereby, the exchange rate in levels introduces an additional stochastic trend into the system. Trivariate cointegration between the interest rates and the exchange rate accounts for the missing stationarity property of the interest differential. We apply the concept to the case of Turkey and Europe, where we can validate the theoretical considerations by multivariate time series techniques.

JEL-Classification: E44, F31, C32
Keywords: Uncovered Interest Rate Parity, Monetary Policy Rules, Cointegration, Vector-Error Correction Model

1 Introduction

The *uncovered interest rate parity* (UIP), the international no arbitrage capital market relationship, is a key building block of many macroeconomic theories. It is prevalently used in macroeconomic models for open economies (see HOLTEMÖLLER (2005), MERLEVEDE ET AL. (2003) and OBSTFELD and ROGOFF (1995)). However, a huge strand of empirical literature has shown that the UIP relationship does not hold in many cases. FARHI and GABAIX (2008) even classify the failure of the UIP as "a major puzzle in international macroeconomics". This deviation from the theoretically implied UIP mechanism is often referred to as the *interest rate parity puzzle* or the *forward premium anomaly* (see for instance LEWIS (1995), ENGEL (1996) and more recently CHRISTIANSEN ET AL. (2010) and ILUT (2010)). A large literature deals with the sketched UIP anomaly and presents various solutions. A commonly applied approach to analyse the UIP puzzle is based on a single equation regression, where the change in the exchange rate is regressed on the interest differential and a constant. In this simplified setting, in order for the UIP relationship to hold, the constant should equal zero and the coefficient of the interest differential must be equal to unity.[1] These estimation approaches are based on the so called *risk-neutral efficient-market hypothesis*. Early studies (eg. FAMA (1984), FROOT and THALER (1990), MCCALLUM (1994)) as well as more recent studies (eg. CHINN and FRANKEL (2002), FRANKEL and POONAWALA (2010)) found parameter estimates significantly deviating from theory. Explanations for this empirical phenomenon vary from market irrationality and sample biases to the existence of currency risk premia (see for instance FROOT and FRANKEL (1989), CAVAGLIA ET AL. (1994) and ENGEL (1996)).

In contrast to these approaches, the present paper focuses on the cointegration property implied by UIP. Cointegration between the two interest rates is a necessary condition in order for the UIP to hold. The absence of bivariate cointegration between domestic and foreign interest rates, however, is a frequently detected empirical failure of the UIP. While for instance JUSELIUS (1995) and JUSELIUS and MACDONALD (2004) relate non-stationarity of interest rate differentials to long-run deviations from purchasing power parity, WOLTERS (2002) links the UIP to the term structure of interest rates so as to revive the cointegration property.

This paper, on the other hand, focuses on the role of monetary policy as a source of non-stationarity of the interest rate differential. In this framework, central banks' reaction on exchange rate fluctuations is the crucial element. Monetary policy is regarded as an intermediary mechanism connecting interest- and exchange rates via their monetary policy reaction function (see for instance MCCALLUM (1994), CHINN and MEREDITH (2004) and WEBER (2010)). Thereby, the exchange rate affects the interest rate differential by introducing an additional stochastic trend into the system.

Here, we consider the European-Turkish case. This constellation is especially interesting with respect to the role of the exchange rate in the conduct of monetary policy. It is

[1] By relaxing the risk-neutrality assumption a constant can reflect a constant risk premium (e.g. see CHINN and MEREDITH (2005)).

reasonable to assume that the Turkish central bank policy is, inter alia, affected by the Euro to Turkish Lira (EUR/TRY) currency exchange rate and the European policy rate. This deliberation can be justified by political reasons in the course of the EU accession process and by growing economic interdependencies between both regions. Moreover, by analysing to what extent the Central Bank of the Republic of Turkey (CBRT) targeted exchange rate developments over a period from 1987 to 2009, CIVCIR and AKCAGLAYAN (2010) showed that exchange rate developments did play a role in Turkish monetary policy. Indeed, we will show empirically that the inclusion of the exchange into the interest rate differential leads to trivariate cointegration between the three rates.

In terms of empirical methodology, central estimations are based on *vector error correction models* (VECMs), where we follow the procedure proposed by JOHANSEN (1995). In contrast to many conventional UIP studies, a system approach is favoured over a single equation approach for reasons of efficiency and due to endogeneity of the exchange rate.

The paper is structured as follows. Section 2 introduces the theoretical concept of the UIP, explains the role of monetary policy and implements the connection to the exchange rate. This is followed by the empirical results with emphasis on the cointegration analysis and the VECM estimation which are of prior importance for the underlying research question. In this section robustness issues are also discussed. Finally, section 4 sums up the main findings and provides concluding remarks.

2 Theoretical Background

On the basis of arbitrage considerations, the UIP states that interest differentials between assets of two different countries but with similar characteristics and equal maturity equal the expected change in the exchange rate. Hence, UIP implies that high-yielding currencies depreciate and low-yielding currencies appreciate so that currency revaluation exactly offsets the interest differential. This economic rationale of the UIP can be formalized as follows:

$$i_{t,m} - i^*_{t,m} = \frac{12}{m}(E_t s_{t+m} - s_t) + \varepsilon_{t,m} \; . \tag{1}$$

This is the logarithmic UIP version, where $i_{t,m}$ denotes the annualised domestic interest rate at time t with m month maturity and $i^*_{t,m}$ denotes the annualised interest rate of the reference country. s_t is the logarithm of the spot exchange rate in terms of domestic currency units per foreign currency unit. E_t represents the conditional expectation operator, where $E_t s_{t+m} \equiv E(s_{t+m}|\Omega_t)$ and Ω_t being the information set available at time t. The factor $\frac{12}{m}$ annualises the expected change in the exchange rate. $\varepsilon_{t,m}$ represents the logarithm of a possible risk premium which allows for the case of imperfect substitutability between different assets and currency risk.

As discussed in the introduction, we focus on the cointegration properties of equation (1). First, assume the exchange rate to be integrated of order one (I(1)), what is in line with a large strand of empirical literature. Then, under rational expectations, the expected change in the exchange rate is stationary. Presupposed that the risk premium $\varepsilon_{t,m}$ follows an I(0) process, in order for the UIP to be valid (and the whole equation being balanced) the interest differential on the left hand side of the equation should be stationary, too. Therefore, provided the case that both interest rates follow I(1) processes (what will be shown below), the domestic and foreign rates should be cointegrated with vector (1,-1). However, as outlined in section 1, the interest differential often does not meet the implied cointegration condition of the UIP.

A popular explanation for deviations of the UIP is the existence of non-stationary risk premia. In contrast, we explain the non-stationarity of the interest differential via central bank behaviour. More precisely, we allow the monetary policy reaction function to be, inter alia, influenced by exchange rate variations. As already mentioned, central bank policy is regarded as an intermediary mechanism connecting exchange- and interest rates. A currency revaluation, for instance, may lead to monetary policy responses in order to counteract rapid changes in exchange rates (see for example McCallum (1994)). Interest rates used in this analysis are short-term rates and, hence, to a huge extend influenced by central bank policy. As shown in equation (1), UIP alone leaves no room for the exchange rate in levels to have an influence on the interest rate differential. However, empirical results in section 3 will justify the reasoning that the exchange rate introduces a stochastic trend into the interest differential. Logically, cointegration between the interest differential and the exchange rate accounts for the frequently detected failure of the implied UIP cointegration condition. In order to account for these mentioned considerations we specify the following stylised policy rules. The policy rule representations serve the purpose of our analysis which is focused on the cointegration property of the UIP rather than the examination of extensive policy rules. Hence, with regard to their specification, the rules presented are formalized in a general manner.

At first, the monetary reaction function for the domestic country is formalized as:

$$\Delta i_{t,m} = F_a(L)\Delta i_{t,m} + \frac{1}{\theta} v_{2t} \ , \tag{2}$$

with $F_a(L) = \sum_{j=1}^{a} f_j L^j$ and $\theta > 0$. Given the case that the Euro area represents the domestic country and Turkey the reference country, it is realistic to assume that the European Central Bank (ECB) policy is not affected by the EUR/TRY exchange rate and, hence, s_t is not included in equation (2). The empirical results of section 3 will support this reasoning. Due to the I(1) property of the data (see section 3.2), equation (2) is given in first differences. Economically, this formulation represents interest rate smoothing which keeps the interest rate from departing too far from its recent values. The inclusion of the lag-polynomial $F_a(L)$ provides some flexibility to this effect. The stationary stochastic process v_{2t} represents other relevant factors like output- and inflation gap. After rearranging and integrating equation (2) the policy rule has the following form:

$$\theta i_{t,m} = \frac{v_{2t}}{1-L} \frac{1}{(1-F_a(L))} \ . \tag{3}$$

Note that $\frac{v_{2t}}{1-L} = \sum_{j=0}^{\infty} v_{2t-j}$ is the stochastic trend of the interest rate and, hence, represents the long-run component of the series.

The policy rule for the foreign country, Turkey in our application, is given as:

$$\Delta i^*_{t,m} = W_p(L)\Delta s_t + B_q(L)\Delta i^*_{t,m} + u_t \ , \tag{4}$$

$$u_t = C_h(L)\Delta v_{1t} + v_{2t} \ , \tag{5}$$

where $W_p(L) = \sum_{j=1}^{p} w_j L^j$ and $B_q(L) = \sum_{j=1}^{q} b_j L^j$. As outlined above, from an economic perspective we expect $W_p(1) < 0$. u_t denotes a stationary stochastic process representing other relevant policy influences, where $C_h(L) = \sum_{j=0}^{h} c_j L^j$. Thereby, the component v_{1t} (representing Turkish policy shocks) is solely related to the Turkish interest rate, whereas the second stochastic component, v_{2t}, is the common stochastic process of the domestic and foreign interest rate, as introduced in equation (2). Hence, v_{2t} represents Euro area policy shocks affecting the conduction of Turkish monetary policy. Thereby, we model dependence of Turkey on the stance of European monetary policy and set the stage for cointegration between the two interest rates and the exchange rate. As for the European policy rule, we account for the existence of smoothing effects. In contrast to the Euro area policy rule, this specification allows the lagged values of the exchange rate to exert influence on central bank's behaviour.[2] The exchange rate is expected to be negatively related to the interest rate. This implies that, for example, a decreasing exchange rate (bear in mind the direct notation) induces the interest rate $i^*_{t,m}$ to rise in order to counteract currency devaluation.

Integration of equation (4) and (5) yields:

$$i^*_{t,m} = W_p(L)s_t + B_q(L)i^*_{t,m} + C_h(L)v_{1t} + \frac{v_{2t}}{1-L}\ , \tag{6}$$

where $\frac{v_{2t}}{1-L}$ is exactly the same stochastic trend as incorporated in the European reaction function. The stochastic process $C_h(L)v_{1t}$, however, remains transitory. Therefore, it matters for short- and medium-run dynamics but is not of importance for the cointegrating relationship.

Due to the common stochastic trend between Turkish and European interest rate, we can substitute equation (3) into equation (6).

$$\underbrace{i^*_{t,m}}_{I(1)} = \underbrace{W_p(L)s_t}_{I(1)} + \underbrace{B_q(L)i^*_{t,m}}_{I(1)} + \underbrace{C_h(L)v_{1t}}_{I(0)} + \underbrace{\theta(1-F_a(L))i_{t,m}}_{I(1)} \tag{7}$$

Equation (7) contains a linear combination of I(1) variables and, hence, theoretically implies trivariate cointegration between the interest rates and the exchange rate. The corresponding long-run relationship can be expressed as follows:

$$\overline{i^*}_{t,m} = \frac{W_p(1)}{1-B_q(1)}\bar{s}_t + \frac{\theta(1-F_a(1))}{1-B_q(1)}\bar{i}_{t,m}\ . \tag{8}$$

Following our theoretical hypotheses the exchange rate enters the equilibrium relation with a negative coefficient since we expect $W_p(1) < 0$ and $(1 - B_q(1)) > 0$ and the interest rate of the domestic country, $\bar{i}_{t,m}$, with a positive parameter since $\theta > 0$ and $(1 - F_a(1)) > 0$. We now take the model to the data in order to empirically validate the derived cointegration property.

[2]Instead of using nominal exchange rates also real exchange rates could be considered (see ENGEL and WEST (2006)). However, we also tested the unit root property of the real exchange rate. It turned out that the real exchange rate also follows an I(1) process and introduces an additional stochastic trend into the system. Hence, in this respect our results do not hinge on the choice between nominal or real exchange rates.

3 Empirical Results

3.1 Data

The interest rates used in this analysis are annualised short-term money market rates. The European interest rate is a combination of the 3 month Frankfurt banks money market rate (from January 1985 to December 1998) and the Euribor 3 month rate (from January 1999 on). This takes into account the leading role of the Bundesbank in pre-European Monetary Union (EMU) monetary policy. For the Turkish interest rate, the Public Time Deposit 3 month rate is used. The interest rates are measured in percentage points. The EUR/TRY exchange rate before January 1999 was constructed through the triangular relation with the TRY/USD and the USD/EUR exchange rate. The EUR/TRY exchange rate is transformed to logarithm and multiplied by 100, so that taking first differences would generate continuously compounded monthly returns in percentages. The series are depicted in Figure 1 to 3.

The sample contains monthly data and goes from January 1985 to March 2010, yielding a sample size of 302 observations. This sample choice covers the Turkish liberalization process of the 1980s, which included removing price controls, subsidies and interest rate ceilings, establishing an interbank money market and the Istanbul Stock Exchange, liberalising foreign trade, relaxing capital controls, encouraging foreign direct investment and expanding the private sector as well as starting to use indirect monetary policy instruments in policy implementation (MACOVEI (2009) and BERUMENT and DINCER (2008)). These reforms are essential conditions in order for the basic UIP assumptions (e.g. no capital-movement restrictions and reduced transaction costs) to hold. All data have been obtained using national sources via Datastream.

Regarding the European interest rate in Figure 1, the series exhibits no clear down- or upward trending behaviour and goes through various peaks and troughs. The steep and sudden fall in the interest rate at the end of 2008 stands out and can clearly be related to the events of the global financial crisis. The Turkish interest rate, in contrast, is considerably higher throughout the whole sample. Especially remarkable are its two peaks in the years 1994 and 2001, the trough during the year 2000 as well as the downward trending behaviour from 2002 on. These characteristic features of the curve mirror incisive events in recent Turkish economic history.

In 1994, Turkey was affected by a severe economic crisis. In the years prior to this event, the Turkish government's fiscal balances deteriorated and its rising public sector borrowing requirement relied heavily on central bank financing. This lack of fiscal discipline and sound debt management led rating agencies to downgrade the Turkish economy. These events finally triggered the financial crisis. As a consequence and due to a lack of foreign exchange reserves of the CBRT, the currency was devalued and monetary- and fiscal policy tightened. The nominal interest rate of government securities even reached 100% (EMIR ET AL. (2000)).

In the late 1990s, the Russian crisis in 1998, general elections as well as two earthquakes in 1999 deteriorated public sector fiscal balances once again. In 2000/2001, Turkey experienced its severest economic crisis to date. In the course of disinflation measures and due to Turkey's fragile banking- and financial sector, interest rates first undershot and then started to skyrocket. These events finally lead to the float of the Turkish lira in February 2001.[3] The downward trending behaviour of the interest rate from 2002 on is indicative for the success of the post-crisis implemented structural reforms and sound macroeconomic policies which helped the Turkish economy to recover swiftly.

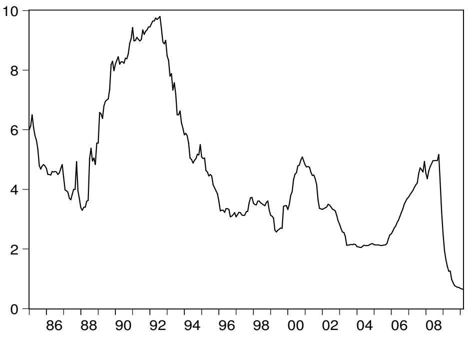

Figure 1: **European 3 month money market rate**

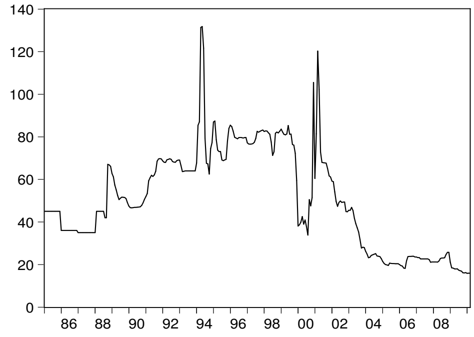

Figure 2: **Turkish 3 month deposit rate**

Figure 3: **European-Turkish exchange rate**

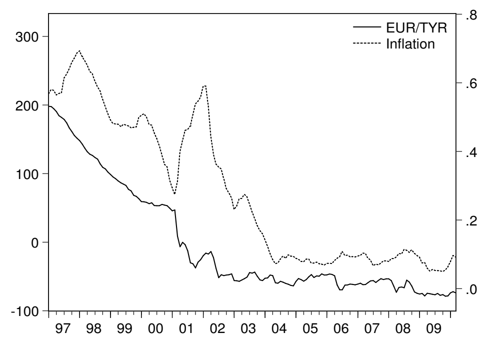

Figure 4: **European-Turkish exchange rate and Turkish inflation rate**

The EUR/TRY exchange rate curve of Figure 3 clearly shows a downward trending behaviour but a clear break of this depreciating process in 2001/2002.[4] As shown in Figure 4, this can be, inter alia, attributed to the moderation of Turkey's persistently high inflation rates that were brought down by central bank measures.[5] As a consequence of

[3]For further details of the Turkish liquidity and currency crisis see for instance ALPER (2001) or MACOVEI (2009).
[4]Later in the econometric modelling we will consider that break.
[5]As the relevant measure for inflation we use the first difference of the log of the Turkish consumer price

the 2000/2001 crisis, the CBRT floated the exchange rate and was granted instrument independence which enabled the bank to set monetary policy autonomously. Its main goal was defined as maintaining price stability. From 2002 to 2005 the CBRT followed an implicit inflation targeting regime and changed to an explicit inflation targeting regime in 2006 (ALPER and HATIPOGLU (2009)). As a result, notoriously high inflation rates declined and the exchange rate finally stabilized from 2002 on.

3.2 Unit Root Test

The theoretical model assumed the variables to be non-stationary. This property shall now be established in order to set the stage for the cointegration analysis. The unit root behaviour of the time series is checked by ADF tests (see DICKEY and FULLER (1979)). Critical values for the null hypothesis of non-stationarity are taken from MACKINNON (1996). Here, as well as in all subsequent models, the lag length is chosen following the Akaike information criterion (AIC) and Lagrange Multiplier (LM) residual autocorrelation tests.

Note that a linear trend would not be meaningful for interest rates and is also not supported by the data. However, for the exchange rate a trend is not a priori ruled out, for instance reflecting nominal depreciation due to inflation differentials. Hence, as far as the levels are concerned the ADF test equation for the interest rates includes a constant, whereas the test equation for the exchange rate additionally incorporates a trend. For the first differences, no deterministics are included in the test equation for the interest rates but a constant is considered for the exchange rate equation. Table 1 summarizes the unit root test results.

Table 1: **ADF tests**

	Levels			First differences		
Variable	$\hat{t}$		l	$\hat{t}$		l
$i_{t,3}$	−2.079	(0.254)	10	−3.313	(0.001)	9
$i^*_{t,3}$	−1.567	(0.498)	5	−9.891	(0.000)	4
s_t	0.923	(0.999)	1	−15.367	(0.000)	0

Note: Test statistics are denoted by $\hat{t}$. l refers to the number of lagged differences and p-values are given in parentheses.

In none of the cases, the null of a unit root can be rejected at the 10% significance level. The first differences, however, are clearly stationary at the 1 % level. Hence, we assume all series integrated of order one. The results remain robust even if we alter the sample size or apply different information criteria for the choice of the lag length.[6] Even when

index (as year on year change). The data are also obtained from Datastream and measured by the right axis in percentage points.

[6]When applying a KPSS test to check for the null of stationarity, the obtained results equal the findings

considering a trend break for the exchange rate does not alter the results.[7] Thus, we focus on the cointegration analysis in the following.

3.3 Cointegration Analysis

Given the results of section 3.2 a multivariate non-stationary model should be applied. We choose the JOHANSEN (1995) cointegration approach in the following. The likelihood ratio test (JOHANSEN (1994, 1995)) is used to test for the number of cointegrating relationships in a VAR. The corresponding test statistic is given by

$$Tr(r_0) = -T \sum_{i=r_o+1}^{n} ln(1-\hat{\lambda}_i) \ , \tag{9}$$

where n denotes the number of endogenous variables and T the sample size. $\hat{\lambda}_i$ represents the i-th largest squared canonical correlation between Δy_t and y_{t-1} from equation (10) of section 3.4, both corrected for the influence of the remaining regressors.

Since the UIP implies cointegration of the interest rates, we test for bivariate cointegration between the European- and Turkish interest rate. The test incorporates a constant restricted to the error correction term as appropriate for interest rates. The test result is depicted in Table 2.

Table 2: **Likelihood ratio trace test for European- and Turkish interest rates**

	Test statistic	p-value	lag length
$H_0 : r_0 = 0$	7.47	0.860	6

As already mentioned, the absence of bivariate cointegration between domestic and foreign interest rates is a frequently detected empirical failure of the UIP. Indeed, the null of no cointegration cannot be rejected at any reasonable significance level. Hence, we follow the theoretical considerations made in section 2 and include the EUR/TRY exchange rate into the system. In this way, cointegration can potentially be restored via a trivariate cointegrating relationship as shown by equation (7).

Figure 3 suggests that a deterministic trend should be considered. Furthermore, since the downward movement came to an end in the latter part of the sample, we further include a trend shift. Technically, this implies for the VECM in (10) of section 3.4 that the deterministic part d_{t-1} inside the cointegrating relationship consists of a linear trend and a shift in the slope of this trend and that the deterministic component D_t outside the long-run relation contains a constant and a shift dummy (see also JOHANSEN ET AL. (2000)).

from the ADF test that all series are integrated of order one. This implies that a lack of statistical power does not stand behind the I(1) result.

[7] We applied the JOHANSEN ET AL. (2000) procedure as explained in the following section in a minimum setup by solely incorporating the exchange rate.

As the trend break point, January 2002 is chosen due to its significance in Turkish monetary policy; see section 3.1. Although the Turkish lira was floated already in February 2001, this process was followed by a considerable depreciation of the exchange rate. In 2002, the CBRT implemented an implicit inflation targeting regime with the short-term interest rate as its main policy instrument. Only from that point on the Turkish exchange rate movement started to stabilize. The results of the trivariate trace tests are depicted in Table 3.

Table 3: **Likelihood ratio trace test for trivariate cointegration between the interest rates and the exchange rate**

	Test statistic	p-value	lag length
$H_0 : r_0 = 0$	85.86	(0.000)	4
$H_0 : r_0 \leq 1$	29.91	(0.210)	4

The results of Table 3 strongly support the theoretical idea of trivariate cointegration between the exchange rate and the two interest rates. The null hypothesis of no cointegration is clearly rejected, whereas the null of one cointegrating relationship cannot be rejected at any reasonable significance level. The results remain unaffected by the use of different information criteria. Thus, we assume one cointegrating relationship between the three series and proceed with the VECM analysis.[8]

3.4 Vector Error Correction Analysis

Given the results of the precedent section that one cointegrating relationship between the series of the system is present, Granger's representation theorem leads to the VECM:

$$\Delta y_t = \alpha(\beta' y_{t-1} + d_{t-1}) + \sum_{i=1}^{q} A_i \Delta y_{t-i} + D_t + u_t \ , \qquad (10)$$

with A_i, $i = 1, 2, ..., q$, capturing the coefficients of the short-run dynamics. y_t is a n-dimensional vector of the endogenous variables $i_{t,3}$ $i^*_{t,3}$ and s_t. u_t denotes a n-dimensional vector of the residuals and D_t as well as d_{t-1} represent vectors including the deterministic components of the system. According to the considerations made in section 3.3, the deterministic components are specified outside and inside the cointegrating relationship. The vector α contains the adjustment coefficients, β is the cointegration vector and, hence, $\beta' y_{t-1}$ represents the stationary linear combination of the variables.

Since the exchange rate is in general an endogenous variable, we estimate the model in a multivariate system. The VECM analysis offers additional useful insights into the empirical validity of our theoretical considerations with regard to the long-run relation and

[8] We also checked if the inclusion of a trend and a trend break alters the test for bivariate cointegration between the interest rates. The established result remained unchanged.

the variables' adjustment to equilibrium deviations. Moreover, it allows distinguishing between short- and long-run dynamics of the system.

First, we estimated the VECM with one cointegrating relation and deterministics as specified in the preceding section. Thereby, the trend and trend shift inside the cointegrating relationship turned out to be not significant, implying that they are orthogonal to the cointegration space. In contrast, the deterministic part outside remained significant. Consequently, we estimate a VECM including constant term (capturing the mean and orthogonal trend) and level shift outside the cointegrating relationship.[9] The lag length is, according to the AIC and the LM test results for no residual autocorrelation (depicted in Table 4), set to three and the standard deviations are given in parentheses. After applying a model reduction procedure[10], equation (11) represents the relevant parts of our VECM estimation:

$$\begin{pmatrix} \Delta i_t \\ \Delta i_t^* \\ \Delta s_t \end{pmatrix} = \begin{pmatrix} 0.000 \\ \underset{(0.057)}{0.279} \\ \underset{(0.013)}{0.131} \end{pmatrix} \left[i_{t-1} - \underset{(0.077)}{0.478} i_{t-1}^* - \underset{(0.006)}{0.024} s_{t-1} \right] + \sum_{j=1}^{3} \widehat{A}_j \begin{pmatrix} \Delta i_{t-j} \\ \Delta i_{t-j}^* \\ \Delta s_{t-j} \end{pmatrix} + \widehat{D}_t + \widehat{u}_t \,. \quad (11)$$

Table 4: **LM test results for null of no residual autocorrelation in a VECM specified with three lags**

lag length	1	2	6	12
p-value	0.840	0.906	0.462	0.264

Several points are important to address. First, the long-run equilibrium can be written as:

$$0.478\overline{i_t^*} = \overline{i_t} - 0.024\overline{s_t} \,. \quad (12)$$

Equation (12) parallels the theoretical relation (8) in that the Turkish interest rate is expected to be positively related to the European interest rate and negatively related to the exchange rate in the long-run. This relationship empirically validates our theoretical reasoning on Turkish central bank policy.

The adjustment coefficient of the European interest rate is insignificant and, thus, the European interest rate weakly exogenous. This outcome was already expected in section 2 (see equation (2)) and appears quite plausible. The ECB is not reacting to the Turkish interest rate or the Turkish exchange rate. On the other hand, the Turkish interest

[9] The trace test including an orthogonal trend and level shift also significantly rejects the $H_0 : r_0 = 0$ with p-value 0.0003 and does not reject the null of $r_0 \leq 1$ at all reasonable significance levels (p-value 0.423). Hence, assuming one cointegrating relationship is unaffected by the exclusion of the insignificant deterministic components.

[10] Starting from the last regressor in the model, the parameter elimination procedure used in this study sequentially deletes these coefficients in the VECM whose elimination reduces the AIC value. Otherwise they are maintained.

rate is significantly reacting to equilibrium deviations and, hence, is affected by the exchange rate as well as by the European interest rate. Since Turkey is a relatively small economy with strong political and economic links towards Europe, these influences are plausible. The adjustment coefficient of the exchange rate proves significant, too. Furthermore, both adjustment coefficients have the correct signs with regard to an adjustment back to equilibrium in case of a deviation. For instance, imagine the Turkish exchange rate to depreciate. That would induce the Turkish interest rate to rise in order to restore equilibrium. This is also a sensible reaction from an economic perspective since monetary policy tightening typically counteracts tendencies of devaluation. The sign of the exchange rate's adjustment coefficient also restores equilibrium and is in line with the UIP relationship. The size of the adjustment parameters is economically justifiable. In one period, an equilibrium deviation of one unit leads to an adjustment of the Turkish interest rate by 0.279 percentage points and of the exchange rate by 0.13%. This shows that the long-run equilibrium acts as an attractor that exerts important influences on interest and exchange rate.

In order to check the robustness of our results, we altered the sample length by excluding the last part of the sample which captures the period of the recent global financial and economic crisis. We took both the event of the Lehman Brothers bankruptcy in September 2008 and the first repercussions in the US housing market in mid-2007 as start date of the crisis. The finding of cointegration and the dimensions of the adjustment coefficients remain unchanged. Therefore, the implications drawn from our analysis do not hinge on the presence of the crisis observations in the sample. The results are also robust with regard to the choice of the lag length following different information criteria.

4 Concluding Summary

This study examined the cointegration property of the UIP relationship. As implied by the UIP, the interest rate differential between domestic and foreign interest rates should be stationary and, hence, the interest rates cointegrated with vector (1,-1). However, the absence of bivariate cointegration between domestic and foreign interest rates is a frequently detected empirical failure of the UIP.

In an European-Turkish context, we explained this anomaly via central bank behaviour and regard monetary policy as a source of non-stationarity of the interest rate differential. Since interest rates used in this analysis are short-term rates, central bank policy is a crucial factor for their determination. Thereby, central banks' reaction function is, inter alia, influenced by exchange rate variations. We showed that this leads to a connection of interest- and exchange rates in levels. If, for instance, a central bank seeks to stabilize or control exchange rate movements, this may lead to a change in the monetary policy rate. Through this process, the exchange rate in levels introduces an additional stochastic trend into the system. The Turkish case is especially interesting with respect to the role of the exchange rate in the conduct of CBRT's monetary policy. Due to political reasons in the course of the EU accession process as well as growing economic linkages between both regions, it is reasonable to assume that Turkish monetary policy is influenced by the EUR/TRY exchange rate.

By applying multivariate time series techniques, our theoretical considerations could be empirically validated. As often found in the literature, the interest rates are not cointegrated. By introducing the exchange rate into the cointegrating relationship, though, a trivariate long-run relationship could be established. The VECM analysis further substantiated the empirical validity of our theoretical considerations. With regard to the long-run equilibrium, higher Turkish interest rates go along with a weaker Turkish currency. Again, this result is in line with our theoretical explications on central bank behaviour. Concerning the adjustment of the variables towards equilibrium, whereas the Turkish interest rate is significantly reacting on equilibrium deviations, the European interest rate, as expected, turned out to be weakly exogenous. Relatively smaller economies like Turkey are more likely to be affected by macroeconomic changes of economically more important economies like the EMU. The exchange rate also significantly restores equilibrium in case of a deviation and is in line with the UIP relationship. Therefore, we conclude that our theoretical model provides an economic rationale of seeming UIP failure which can be supported by the data. This bears the potential to apply our approach to further country combinations. Moreover, introducing the underlying mechanism into a structural modelling context appears as a valuable direction for future research.

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