

PANEL COINTEGRATION ANALYSIS TO  
EXCHANGE RATE DETERMINATION:  
MONETARY MODEL VERSUS TAYLOR RULE MODEL

A Master's Thesis

by

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January 2009



To My Sister

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The Institute of Economics and Social Sciences  
of  
Bilkent University

by

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In Partial Fulfillment of the Requirements for the Degree of  
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in

THE DEPARTMENT OF  
ECONOMICS  
BİLKENT UNIVERSITY  
ANKARA

January 2009

I certify that I have read this thesis and have found that it is fully adequate, in scope and in quality, as a thesis for the degree of Master of Arts in Economics.

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## **ABSTRACT**

### **PANEL COINTEGRATION ANALYSIS TO EXCHANGE RATE DETERMINATION: MONETARY MODEL VERSUS TAYLOR RULE MODEL**

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This thesis examines the validity of the monetary model and the Taylor-rule model in determining exchange rates in the long run. The monetary model and the Taylor-rule model are tested using the US dollar exchange rates over 1980:01-2007:04 periods for 13 industrialized countries. Johansen Fisher Panel cointegration technique provides evidence that there exist a unique cointegration relationship between the nominal exchange rates and a set of fundamentals implied by the monetary model and the Taylor rule model. The cointegrating coefficient estimates for the monetary model and the Taylor rule model are found by using panel dynamic ordinary least square (DOLS) estimator. The estimation results show that the effects of the monetary and the Taylor rule fundamentals on exchange rates are not the same as what the theory suggests. Overall, the findings of this thesis imply that there is no support for the monetary model and there is little support for Taylor-rule model in explaining exchange rates.

**Keywords:** Exchange Rates, Monetary Model, Taylor Rule, Panel Unit Root Tests, Panel Cointegration

## ÖZET

### DÖVİZ KURUNUN BELİRLENMESİNDE PANEL EŞBÜTÜNLEŞME ANALİZİ: PARASAL MODEL İLE TAYLOR KURALI MODELİ'NİN KARŞILAŞTIRILMASI

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Bu tez, uzun dönemde döviz kurlarının belirlenmesinde parasal model ile Taylor kuralı modelinin geçerliliğini araştırmaktadır. Parasal model ve Taylor kuralı modeli Amerikan doları döviz kuru kullanılarak 1980: 01 – 2007: 04 dönemleri arasında 13 sanayileşmiş ülke için test edilmiştir. Johansen Fisher panel eşbütünleşme tekniği nominal döviz kurları ile parasal model ve Taylor kuralı modelinin temel değişkenleri arasında tek bir eşbütünleşme ilişkisi olduğunu kanıtlamaktadır. Parasal model ile Taylor kuralı modeli için eşbütünleşme katsayı tahminleri panel dinamik en küçük kareler (DEKK) tahminleyicisi kullanılarak bulunmuştur. Tahmin sonuçları parasal model ile Taylor kuralı modelinin temel değişkenlerinin döviz kurları üzerindeki etkisinin teorinin öne sürdüğü etki ile aynı olmadığını göstermektedir. Genel itibariyle bu tezin bulguları parasal modelin döviz kurunu açıklama gücüne yönelik kanıt bulunmadığını ve Taylor kuralı modelinin döviz kurunu açıklama gücüne yönelik az kanıt bulunduğunu işaret eder.

Anahtar Kelimeler: Döviz Kurları, Parasal Model, Taylor Kuralı, Panel Birim Kök Testleri, Panel Eşbütünleşme

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## **CHAPTER I**

### **INTRODUCTION**

During the past two decades, researchers have tried to test the empirical validity of Frenkel's (1976) flexible-price monetary model of the exchange rate by using different samples and estimation techniques. The earlier studies searched a long-run relationship between exchange rates and the monetary fundamentals by employing an Engle Granger two step procedure on the time series of the individual countries. Among these studies, particularly Boothe and Glassman (1987), Baillie and Selover (1987), and McNown and Wallace (1989) could not find much of an evidence for a cointegration relationship between exchange rate and its main determinants suggested by the monetary model of Frenkel.

The failure of the monetary model of exchange rate by using Engle Granger technique has led many researchers to employ another methodology namely, Johansen's multivariate cointegration technique. MacDonald and Taylor (1993, 1994), McNown and Wallace (1994), Diamandis et al. (1996) find evidence in favor of the monetary model in determining exchange rates by using

Johansen's multivariate cointegration methodology. Cushman (2000) also finds evidence in favor of cointegration between the US dollar-Canadian dollar exchange rate and a set of monetary fundamentals by using Johansen's methodology; however, the estimated cointegrating coefficients differ widely from those suggested by the monetary model. Cushman (2000) therefore concludes that there is no support for the monetary model in US-Canadian data.

With the recent developments in panel unit roots and panel cointegration analysis, these new techniques are used to check the validity of the monetary model of exchange rate in the long-run. Among the studies which utilize panel data techniques, particularly Groen (2000, 2005) and Mark and Sul (2001), and Rapach and Wohar (2004) provide noteworthy panel results. Their results suggest that the monetary model might explain nominal exchange rate trends during post-Bretton Woods float.

All of the aforementioned studies using both time series and panel data techniques which are testing the validity of the monetary model assume that central banks are using money supply as a monetary policy instrument. However, especially in industrialized countries modern central banks have used short-term interest rates as the single monetary policy tool. In fact, Clarida et al. (1998) observe that, since 1980's, the central banks of Germany, Japan and U.S. have pursued inflation targeting meaning that in response to a rise in expected inflation relative to target, each central bank raises nominal rates sufficiently to push up real rates.

On the basis of the changes in monetary policy instruments of the central banks, Clarida et al. (1998) estimate monetary policy reaction functions assuming that central banks set the nominal short term interest rates according to a simple interest rate rule proposed by Taylor (1993). They find that the coefficient of real exchange rate in the interest rate rule is statistically significant for Germany and Japan.

The findings of Clarida et al. (1998) gave rise to a new strand of literature, notably Engel and West (2005, 2006), Mark (2005), and Molodtsova and Papell (2008), which examines the linkage between the exchange rates and a set of fundamentals that arise when central banks set the interest rate according to the Taylor-rule. Engel and West (2006) and Mark (2005) finds evidence Taylor-rule fundamentals provide a plausible framework for understanding real dollar-DM exchange rate dynamics. Molodtsova and Papell (2008) assess the out-of-sample performance of the monetary and the Taylor rule models and provide the evidence of predictability is much stronger with Taylor rule model than with the monetary model at short horizon. Among them, Engel and West (2005) failed to find a cointegration relationship between exchange rates and a set of fundamentals implied by the monetary and the Taylor rule models.

Earlier studies which use time series and panel-based frameworks generally investigate whether the exchange rates are cointegrated with monetary fundamentals. In the time series literature, only the study of Engel and West (2005) examines the linkage between exchange rates and Taylor rule fundamentals using cointegration analysis. In other words, the role of Taylor-

rule fundamentals in explaining exchange rates with panel cointegration techniques has not been studied previously. Hence, this thesis intends to contribute to the existing literature by using a panel cointegration technique to examine the linkage between the exchange rates and the Taylor-rule fundamentals.

This thesis is also motivated from the fact that combination of cross-sectional and time-series information in the form of a panel data set can greatly increase the power of the unit root and cointegration tests. We aimed to investigate whether the nominal exchange rates are cointegrated with the monetary and the Taylor rule fundamentals in a panel based framework.

All of the previous studies on the monetary fundamentals utilize panel cointegration tests based on Engle-Granger (1987) framework which depends testing the stationarity of the residuals from a levels regression. Unlike the previous literature this thesis uses the panel cointegration analysis by employing Johansen Fisher panel cointegration technique which is originated from Johansen's multivariate cointegration methodology.

In this thesis, Johansen Fisher panel cointegration technique is applied to check for the validity of the monetary and the Taylor-rule models of exchange rate in the long run. The monetary and the Taylor-rule models are tested using the US dollar exchange rates over 1980:01-2007:04 periods for 13 industrialized countries. In order to find support for the monetary and the Taylor-rule models, two pieces of evidence are needed. The first one is to find that the US dollar exchange rate is cointegrated with the monetary and the Taylor-rule

fundamentals. The second one is to obtain estimates of the cointegrating coefficients relating the US dollar exchange rate to a set of fundamentals which agree with the values suggested by the monetary and the Taylor-rule models.

The rest of the paper is organized as follows: Section 2 presents the previous studies in the literature. Section 3 gives the theoretical background for the monetary and the Taylor-rule models. The methodology is explained in Section 4. Empirical analysis including data sources, panel unit root and panel cointegration tests, and the estimation results are presented in Section 5. Finally, Section 6 gives the concluding remarks.

## **CHAPTER II**

### **LITERATURE REVIEW**

The roots of the monetary approach in determining the exchange rate movements go back to the early 1970s. Frenkel (1976) deals with the determinants of the exchange rate and develops a monetary view of exchange rate determination. After stating the building blocks of the monetary model, which are purchasing power parity (PPP), money demand function and expectations, he provides empirical evidence for the monetary approach to exchange rate determination using the German hyperinflation case during 1920-1923. On the basis of Frenkel's findings, many researchers have started to investigate the monetary approach to exchange rate determination by using time series analysis for different countries and different exchange rates.

Using Engle-Granger cointegration methodology, Boothe and Glassman (1987), Baillie and Selover (1987) and McNown and Wallace (1989) could not find much of an evidence for a cointegration relationship between exchange rate and its main determinants suggested by the monetary approach. On the other

hand, several studies by MacDonald and Taylor (1993, 1994), McNown and Wallace (1994), Diamandis et al. (1996) employ Johansen's multivariate cointegration technique and their results show that exchange rates are cointegrated with monetary fundamentals in the long run. Although MacDonald and Taylor (1993, 1994) and Diamandis et al. (1996) find cointegrating coefficient estimates suggested by the monetary model, McNown and Wallace (1994) provide coefficients estimates which are not consistent with the monetary model. Cushman (2000) also finds that the US dollar-Canadian dollar exchange rate is cointegrated with a set of monetary fundamentals by using Johansen's methodology; however, the estimated cointegrating coefficients differ widely from those suggested by the monetary model. Therefore, Cushman (2000) concludes that there is no support for the monetary model in US-Canadian data.

With the developments of panel unit roots and panel cointegration analysis, researchers have started to use these techniques to find empirical evidence of monetary models in determining exchange rate movements for different set of countries. Two recent studies by Groen (2000) and Mark and Sul (2001) follow the PPP literature and test the monetary model using panels of post-Bretton Woods data. Groen (2000) considers a panel of US dollar nominal exchange rate, relative money supply, and relative real output level data for 14 industrialized countries covering the period 1973:1-1994:4. He finds that nominal exchange rates are cointegrated with relative money supplies and relative output levels and panel cointegration coefficient estimates that are

reasonably consistent with the monetary model for his full panel and three sub-panels (G10, G7, and EMS).

Mark and Sul (2001) employ a panel of US dollar nominal exchange rate, relative money supply, and relative real output data for 18 countries spanning 1973:1-1997:1. Their results suggest that US dollar exchange rate is cointegrated with monetary fundamentals and that the monetary fundamentals contain significant predictive power for future exchange rate movements. They also find evidence of cointegration using Swiss franc or Japanese yen as the numeraire currency.

Rapach and Wohar (2004) show how poorly the monetary model performs on a country-by-country basis for US dollar exchange rates over the post-Bretton Woods period for a large number of industrialized countries. In addition, using panel analysis they find that panel cointegration tests largely indicate the existence of a long run relationship between nominal exchange rates and monetary fundamentals. Crespo-Cuaresma et al. (2005) use a panel data set for six Central and Eastern European countries to estimate the monetary exchange rate model with panel cointegration methods including the Pooled Mean Group estimator, the Fully Modified Least Square estimator and the Dynamic Least Square estimator. Their findings are in line with the existence of a long run relationship between nominal exchange rates and the monetary fundamentals.

Groen (2005) investigates both the in-sample as well as the out-of-sample fit of monetary exchange rate model in order to assess whether the Euro

exchange rates of Canada, Japan, and the United States have a long-run link with monetary fundamentals. He finds that the aforementioned exchange rates are consistent with monetary exchange rate model based on a common long-run relationship, albeit with a long run impact of relative income that is higher than predicted by the theory. He also provides the evidence that out-of-sample forecasting evaluations based on long-run monetary model are superior to both random walk-based forecasts and standard cointegrated VAR model-based forecasts, especially at horizons of 2 to 4 years.

Traditional exchange rate models assume that central banks are using money supply as an instrument or target variable to implement monetary policy, where money supply is exogenously determined. However, since 1980s, monetary policy instrument used by modern central banks is short term interest rates rather than money supply. In Taylor (1993) original formulation, the rule assumes that the Fed sets the nominal interest rate based on the current inflation rate, the inflation gap-the difference between inflation and the target inflation rate, the output gap-the difference between GDP and potential GDP, and the equilibrium real interest rate. Clarida et al. (1998) estimate monetary policy reaction functions for the US, Germany and Japan. They found the real exchange rate entered an interest rate rule for Germany and Japan with a coefficient that is statistically significant, albeit small.

The literature on exchange rate models with Taylor rule fundamentals is relatively new. Recent papers by Mark (2005), Engel and West (2005, 2006), and Engel et al. (2007), and Molodtsova and Papell (2008), investigate some of

the empirical implications for exchange rates if central banks follow Taylor rules for setting interest rates. Engel and West (2005) find that fundamental variables such as relative money supply, outputs, inflation and interest rates provide little help in predicting changes in floating exchange rates. In contrast, their Granger causality results show that exchange rates should be useful in forecasting future economic variables such as money, income, prices and interest rates.

Mark (2005) examines the implications of Taylor-rule fundamentals for the Deutsche mark-dollar exchange rate determination in an environment where market participants are ignorant of the numerical values of the model's coefficients but attempt to acquire that information using least-squares learning rules. That is to say, he assumes that Taylor rule coefficient values are changing over time. Using quarterly data from 1976 to 2003, he finds evidence that this simple learning environment in the policy rule provides a plausible framework for understanding real dollar-DM exchange rate dynamics.

Engel and West (2006) also use Deutsche mark-dollar exchange rate to test the Taylor rule fundamentals covering the period of 1979-1998. First, they borrow the parameters of Taylor-rule from the study of Clarida et al. (1998). Then they estimate VAR for the variables, namely expected inflation, output gap and interest rates to construct the expected value of the fundamentals. Finally, they use the correlation between the model based and actual real exchange rate as a measure, which is found to be approximately 30 percent.

Engel et al. (2007) emphasize the point that “beating a random walk” in forecasting is too strong a criterion for accepting an exchange rate model. They propose a number of alternative ways to evaluate these models. First they examine in-sample fit by emphasizing the importance of the monetary policy rule and its affects on expectations. Then they present estimates of exchange-rate models in which expected present values of fundamentals are calculated from survey forecasts. Finally, they show that out-of-sample forecasting power of models can be increased by focusing on panel estimation and long-horizon forecasts.

Molodtsova and Papell (2008) also use Taylor rule fundamentals for exchange rate determination. They evaluate the performance of alternative exchange rate models using Clark and West (2006, 2007) inference procedures and assess the out-of-sample performance of the models at 1 to 36 month horizons for a set of 12 currencies over the post-Bretton Woods float. They provide the evidence of predictability is much stronger with Taylor rule models than with conventional models at short horizon, particularly at the one-month-ahead horizon for 8 out of 12 currencies. However, they do not find any statistical evidence of increased predictability of the models relative to a random walk at long horizons.

## CHAPTER III

### THEORETICAL BACKGROUND

#### 3.1. The Monetary Model

The early flexible-price monetary model (Frenkel, 1976) relies on the twin assumptions of purchasing power parity (PPP) and the existence of stable money demand functions for the domestic and foreign economies. The first building block of this model, PPP (absolute), states that exchange rates should equalize the national price levels of different countries in terms of a common currency.

$$s_t = p_t - p_t^* \quad (1)$$

where  $s_t$  is the logarithm of the nominal exchange rate expressed in units of home currency per dollar,  $p_t$  and  $p_t^*$  are the domestic and foreign (US) price levels, respectively.

The second assumption of the model is the stable money demand function in both countries. Monetary equilibria in the home and foreign country respectively are given by:

$$m_t - p_t = b_1 y_t - b_2 i_t, \quad (2)$$

$$m_t^* - p_t^* = b_1 y_t^* - b_2 i_t^* \quad (3)$$

where  $m_t$  is the logarithm of the domestic money supply,  $p_t$  is the logarithm of the domestic price level,  $y_t$  is the logarithm of the domestic income,  $i_t$  is the domestic interest rate, and asterisks denote a foreign country variable. Note that income elasticity denoted as  $b_1 > 0$  and interest semi-elasticity denoted as  $b_2 > 0$  are assumed to be common across countries. Solving for  $p_t$  and  $p_t^*$  in (2) and (3) respectively and plugging into (1) yields the basic exchange rate equation determined by monetary model:

$$s_t = (m_t - m_t^*) - b_1 (y_t - y_t^*) + b_2 (i_t - i_t^*) \quad (4)$$

According to equation (4), an increase in domestic (foreign) money supply will lead the domestic currency to depreciate (appreciate). An increase in domestic (foreign) income will raise the money demand, causing the domestic currency to appreciate (depreciate). Finally, an increase in the domestic (foreign) interest rate will result in depreciation (appreciation) of the exchange rate via a reduction of the demand for money.

A further assumption underlying the monetary model is that domestic and foreign bonds are perfect substitutes so that uncovered interest parity (UIP) holds. UIP suggests that the difference between the domestic and foreign interest rate is just equal to the expected rate of depreciation of the domestic currency:

$$i_t - i_t^* = E_t(s_{t+1}) - s_t \quad (5)$$

where  $E_t(s_{t+1})$  denotes the expectation of the next period's level of the exchange rate conditional on information available in period  $t$ . Then by combining (4) and (5), one reaches an equation as following:

$$s_t = \frac{1}{1+b_2} [(m_t - m_t^*) - b_1(y_t - y_t^*)] + \frac{b_2}{1+b_2} E_t(s_{t+1}) \quad (6)$$

For simplicity set,  $(m_t - m_t^*) - b_1(y_t - y_t^*) = x_t$ . Under rational expectations, by iterating forward, it is easy to show that (6) can be expressed in the forward solution form as following:<sup>1</sup>

$$s_t = (1+b_2)^{-1} \sum_{i=0}^{\infty} [b_2/(1+b_2)]^i E_t(x_{t+i}) \quad (7)$$

where the transversality condition  $\lim_{i \rightarrow \infty} [b_2/(1+b_2)]^i E_t(s_{t+i}) = 0$  has been imposed.<sup>2</sup>

As outlined in Macdonald and Taylor (1993), the exchange rate should be cointegrated with the variables contained in  $x_t$ . This is illustrated by subtracting  $x_t$  from both sides of equation (7) and by rearranging the terms to obtain<sup>3</sup>:

$$s_t - x_t = \sum_{i=1}^{\infty} [b_2/(1+b_2)]^i E_t(\Delta x_{t+i}) \quad (8)$$

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<sup>1</sup> See appendix for detailed solution.

<sup>2</sup> The case where the transversality condition does not hold has examined by MacDonald and Taylor (1993) in detail.

<sup>3</sup> See Appendix for detailed solution.

Now, if the variables entering the  $x_t$  expression are first-difference stationary, I(1), then right hand side of equation (8) must also be stationary. If  $s_t$  is also an I(1) series, the exchange rate must be cointegrated with the variables  $m_t - m_t^*$  and  $y_t - y_t^*$ .<sup>4</sup> So, in the empirical modeling, the following long-run nominal exchange rate equation is used:

$$s_t = \beta_0 + \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + u_t \quad (9)$$

$t = 1, \dots, T$ , where according to theory  $\beta_1 = 1$  and  $\beta_2 < 0$ .

Based on the theoretical findings of the monetary model, panel cointegration tests are employed. It provides evidence for the existence of a long run relationship among the nominal exchange rate, relative money supplies and relative income levels across countries.

### 3.2. The Taylor Rule Model

This subsection examines the linkage between the exchange rates and a set of fundamentals that arise when central banks set the interest rate according to a Taylor rule in a two country model. Let  $\pi_t$  denote the inflation rate in terms of deviation from its target level at time  $t$ , and  $y_t$  be the output gap or deviation of log output from trend at time  $t$ . Following Engel and West (2005, 2006) and Engel et al. (2007), the monetary rules in the foreign and home countries can be described as<sup>5</sup>:

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<sup>4</sup> Under rational expectations forecasting errors are stationary. See Taylor (1991).

<sup>5</sup> Constant terms are omitted for convenience.

$$i_t^* = \gamma_\pi E_t \pi_{t+1}^* + \gamma_y y_t^* + u_{mt}^* \quad (10)$$

$$i_t = \gamma_q (s_t - \bar{s}_t) + \gamma_\pi E_t \pi_{t+1} + \gamma_y y_t + u_{mt} \quad (11)$$

where  $\gamma_\pi > 1$ ,  $\gamma_y > 0$ , and  $\gamma_q > 0$ ,  $i_t$  is domestic interest rate,  $s_t$  is the logarithm of the nominal exchange rate expressed in units of home currency per dollar,  $\bar{s}_t$  is a target for exchange rate,  $E_t$  denotes mathematical expectations conditional on a period  $t$  information set,  $E_t \pi_{t+1}$  denotes deviation of expected inflation from the central bank's target in the domestic country,  $y_t$  is domestic output gap,  $u_{mt}$  error term in the monetary policy of domestic country and asterisks denote a foreign variable (US).

Equation (10) is a standard Taylor rule and equation (11) is a Taylor rule with nominal exchange rate and its target level included. Following Engel and West (2005, 2006) and Engel et al. (2007), it is assumed that two countries have the same monetary policy parameters, namely  $\gamma_\pi$  and  $\gamma_y$ . Equation (10) for the foreign country does not include the nominal exchange rate and its target because there is no evidence that the U.S. has adopted an exchange rate target<sup>6</sup>. Much of the Taylor rule literature puts expected inflation in the monetary rule. It is highly possible that when choosing the target interest rates,  $i_t$  and  $i_t^*$ , the central banks may not have direct information about the current values of the price level. Adding expected inflation allows for this possibility<sup>7</sup>.

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<sup>6</sup> See Molodtsova and Papell (2007).

<sup>7</sup> See Engel and West (2005) for detailed discussion.

To simplify the model, it is assumed that the monetary authorities set the target level of exchange rate to make PPP hold<sup>8</sup>, that is:

$$\bar{s}_t = p_t - p_t^* \quad (12)$$

Since  $s_t$  is expressed in units of home currency per dollar and  $\gamma_q > 0$ , the rule (11) means that home country is assumed to raise interest rates when the currency is depreciated relative to the target. The most important reference can be Clarida et al. (1998), who find that the coefficient of real exchange rate in Taylor rule,  $\hat{\gamma}_q$ , is statistically significant and the estimated values for Germany and Japan are equal to 0.05 and 0.09, respectively.

Subtracting the foreign from the domestic monetary rule, it is obtained

$$i_t - i_t^* = \gamma_q (s_t - \bar{s}_t) + \gamma_\pi [E_t \pi_{t+1} - E_t \pi_{t+1}^*] + \gamma_y (y_t - y_t^*) + u_{mt} - u_{mt}^* \quad (13)$$

Next, write uncovered interest rate parity as

$$i_t - i_t^* = E_t(s_{t+1}) - s_t \quad (14)$$

Now plug uncovered interest rate parity equation into both sides of equation (12) and use the definition of the target level of exchange rate to obtain:

$$s_t = \gamma_q (i_t - i_t^*) + \gamma_q (p_t - p_t^*) - \gamma_\pi (E_t \pi_{t+1} - E_t \pi_{t+1}^*) - \gamma_y (y_t - y_t^*) - u_{mt} + u_{mt}^* + (1 - \gamma_q) E_t(s_{t+1}) \quad (15)$$

Equation (15) is the form of the expected discounted present value models where the discount factor is given by  $(1 - \gamma_q)$ . The observed fundamental in the formulation is given by  $(i_t - i_t^*) + (p_t - p_t^*)$ .

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<sup>8</sup> In Cho and West (2003), target level of exchange rate follows an unobserved random walk.

Following Engel and West (2005), the remaining variables in equation (15) are treated as unobserved. Hence, the interest rate and the price level differences as the Taylor rule fundamentals are used in order to estimate the nominal exchange rate based on the following equation:

$$s_t = \alpha_0 + \alpha_1(i_t - i_t^*) + \alpha_2(p_t - p_t^*) + v_t \quad (16)$$

$t = 1, \dots, T$ , where  $\alpha_1 = \alpha_2 > 0$  according to previous studies.

Based on the theoretical findings of the Taylor-rule model, this thesis tests whether the nominal exchange rates are cointegrated with the relative price levels and the relative interest rates across countries.

## **CHAPTER IV**

### **METHODOLOGY**

This section describes a panel-based framework to conduct panel cointegration analysis for the panel version of (9) and (16). Since a prerequisite for cointegration analysis is that all variables are nonstationary, several panel unit root testing methodologies are employed to determine the order of integration of all variables under study. Traditionally, DF (Dickey-Fuller) or ADF (Augmented Dickey Fuller) tests have been used to test for the presence of unit roots in univariate time series data. In recent years, a number of investigators, notably Levin, Lin and Chu (2002), Hadri (2000), Im, Pesaran and Shin (2003) Maddala and Wu (1999) have developed panel-based unit root tests that are similar to tests applied to individual series. According to Baltagi and Kao (2005), adding the cross sectional dimension to unit root tests can increase the power of the tests due to the information in the time series is enhanced by that contained in the cross-section data. Moreover, in contrast to individual unit

root tests with complicated limiting distributions, panel unit root test statistics have normal limiting distributions<sup>9</sup>. Panel unit root tests developed by Levin et al. (2002), Im et al. (2003), Maddala and Wu (1999), and Hadri (2000) can be summarized as follows:

Levin et al. (2002) uses Augmented Dickey Fuller (ADF) specification by considering the following three models:

$$\Delta y_{it} = \rho y_{i,t-1} + \sum_{L=1}^{p_i} \theta_{iL} \Delta y_{it-L} + \alpha_{mi} d_{mt} + \varepsilon_{it} \quad (17)$$

where  $m=1, 2, 3$ ,  $d_{mt}$  indicates the vector of deterministic variables such as intercepts and time trends and  $\alpha_{mi}$  denotes the corresponding vector of coefficients for model  $m=1, 2, 3$ . In particular,  $d_{1t} = \emptyset$  (the empty set),  $d_{2t} = \{1\}$  and  $d_{3t} = \{1, t\}$ . Hence, LLC test includes fixed effects and individual time trends for each country and it allows the lag order,  $p_i$ , be different for individual cross- section units. The null hypothesis of LLC test is that each individual time series contains a unit root against the alternative that each time series is stationary, i.e.,  $H_0 : \rho = 0$  and  $H_1 : \rho < 0$ .

The major drawback of the LLC test is that it restricts  $\rho$  to be homogeneous across all  $i$ . As Maddala (1999) pointed out, the null may be fine for testing convergence in growth among countries, but the alternative restricts every country to converge at the same rate. Im et al. (2003) (IPS) allow for heterogeneous coefficient of  $y_{it-1}$  in equation (17) and propose an alternative

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<sup>9</sup> See Baltagi and Kao (2005).

testing procedure based on averaging individual unit root test statistics.<sup>10</sup> The null hypothesis is that each series in the panel contains a unit root, i.e.,  $H_0 : \rho_i = 0$  for all  $i$  and the alternative hypothesis allows for some (but not all) of the individual series to have unit roots, i.e.,

$$H_1 : \begin{cases} \rho_i < 0 & \text{for } i = 1, 2, \dots, N_1 \\ \rho_i = 0 & \text{for } i = N_1 + 1, \dots, N \end{cases}$$

where  $N$  is the total number of individual cross-sections units. IPS differs from LLC because all the series in the alternative hypothesis are stationary processes in LLC, while in IPS some series can still be nonstationary in the alternative hypothesis.

An alternative approach to panel unit root tests uses Fisher's (1932) results to derive tests that combine the  $p$ -values from individual unit root tests. Maddala and Wu (1999) propose a Fisher-type test which combines the  $p$ -values from unit root tests for each cross-section  $i$  to test for unit root in panel data. Maddala and Wu (1999) argue that both IPS and Fisher test relax the restrictive assumption of the LLC test that  $\rho_i$  is the same under alternative. That is to say, the null and the alternative hypothesis of the Fisher Augmented Dickey Fuller (ADF) unit root test is the same as for the IPS unit root test. Both the IPS and Fisher tests combine information based on individual unit root tests while IPS requires a balanced panel, Fisher tests can be used to test unbalanced panels.

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<sup>10</sup> Note that IPS always contains intercepts only or intercepts and linear trends. Thus in equation (17)  $m=1$  is not the case for IPS test.

Also, the Fisher test can use different lag lengths in the individual ADF regressions and can be applied to any other unit root tests.

All the panel unit root test mentioned above assume that under the null hypothesis, there is a unit root. In contrast to these tests, Hadri (2000) derives a residual-based Lagrange multiplier (LM) test where the null hypothesis is that there is no unit root in any of the series in the panel against the alternative of a unit root in the panel. This test is also known as a general form of the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test from time series to panel data.<sup>11</sup>

If the presence of a unit root is detected in the variables, then it is necessary to check for the presence of a cointegrating relationship among the variables. There are two types of panel cointegration tests in the literature. The first is similar to the Engle and Granger (1987) framework which includes testing the stationarity of the residuals from a levels regression. The second panel cointegration test is based on multivariate cointegration technique proposed by Johansen (1988). However, panel techniques may be better in detecting cointegration relationships since a pooled levels regression combines cross-sectional and time series information in data when estimating cointegrating coefficients.

Pedroni (1999, 2004) and Kao (1999) extend the Engle-Granger (1987) cointegration test. Kao (1999) proposes Dickey Fuller (DF) and Augmented

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<sup>11</sup> Since EViews6 can only perform homogenous type of Hadri (2000), this test was not used in the empirical part of the thesis.

Dickey Fuller (ADF)-type unit root tests. The DF type test from Kao follows the following model:

$$y_{it} = \alpha_i + \beta x_{it} + e_{it} \quad (18)$$

$$y_{it} = y_{it-1} + u_{it} \quad (19)$$

$$x_{it} = x_{it-1} + \varepsilon_{it} \quad (20)$$

where  $i = 1, \dots, N$  and  $t = 1, \dots, T$ . As both  $y_{it}$  and  $x_{it}$  are random walks, it follows that under the null hypothesis of no cointegration, the residual series,  $e_{it}$ , should be nonstationary. The model has varying intercepts across the cross-section observations, the fixed effects specification, and common slopes across  $i$ . With this model, the DF test can be calculated from the estimated residuals as:

$$\hat{e}_{it} = \rho \hat{e}_{it-1} + v_{it} \quad (21)$$

where  $\hat{e}_{it}$  is the estimated residual of equation (18)

The ADF type test from Kao is based on the estimated residuals of the following equation:

$$\hat{e}_{it} = \rho \hat{e}_{it-1} + \sum_{j=1}^p \gamma_j \Delta \hat{e}_{it-j} + v_{itp} \quad (22)$$

where  $\hat{e}_{it}$  is the estimated residual of equation (18) and  $p$  denotes number of the lags in ADF specification. To test whether  $x_{it}$  and  $y_{it}$  are cointegrated based on DF or ADF test statistics, the null and the alternative hypotheses can be written as  $H_0 : \rho = 1$  and  $H_1 : \rho < 1$ , respectively.

Pedroni (1999, 2004) propose several tests for the null hypothesis of cointegration in panel data model that allows for considerable heterogeneity. Pedroni test differs from Kao test in the sense that it incorporates heterogeneous intercepts and trend coefficients across cross-sections to equation (18) and it assumes  $\rho$  to be heterogeneous across cross-sections in equation (21). Pedroni test constructs four panel statistics and three group panel statistics to test the null hypothesis of no cointegration against the alternative hypothesis of cointegration. In the case of panel statistics,  $\rho$  is assumed to be the same across all the cross sections, thus the null and alternative hypotheses are  $H_0 : \rho_i = 1$  for all  $i$  and  $H_1 : \rho_i = \rho < 1$ , respectively. In the case of group panel statistics  $\rho$  is allowed to vary over the cross sections, thus the null and alternative hypotheses are  $H_0 : \rho_i = 1$  for all  $i$  and  $H_1 : \rho_i < 1$  for at least one  $i$ .

Maddala and Wu (1999) use Fisher-type test to propose an alternative approach to testing for cointegration in panel data by combining tests from individual cross-sections to obtain a test statistic for the full panel. Based on the results of Maddala and Wu (1999), this thesis applies to Johansen Fisher Panel Cointegration test combining individual Johansen's cointegration trace tests and maximum eigenvalue tests. In Johansen's multivariate cointegration technique, Trace Statistic tests for at most  $r$  cointegrating vectors among a system of  $N > r$  time series, and the Maximal Eigenvalue Statistic tests for exactly  $r$  cointegrating vectors against the alternative hypothesis of  $r+1$  cointegrating vectors.

Johansen's multivariate cointegration technique in a panel framework consists of two steps. The first one is the estimation of a panel Vector Autoregression (VAR) model and the second one is the determination of whether an intercept and/or a trend enter cointegration analysis. In general, five distinct models can be considered although the first and the fifth model are not likely to happen and are also implausible in terms of economic theory.<sup>12</sup> Therefore, the problem reduces to a choice of one of the three remaining models (Model 2, 3, 4)<sup>13</sup>.

Johansen (1992) suggests applying the Pantula principle in order to decide which model should be used. The Pantula principle involves the estimation of all three models and the presentation of the results from the most restrictive hypothesis (i.e.  $r = \text{number of cointegrating relations} = 0$  and model 2) through the least restrictive hypothesis (i.e.  $r = \text{number of variables entering the VAR} - 1 = n - 1$  and model 4). The model selection procedure then includes moving from the most restrictive model, at each stage comparing the trace statistic to its critical value, stopping only when it is concluded for the first time that the null hypothesis can not be rejected.<sup>14</sup>

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<sup>12</sup> Model 1: No intercept or trend in CE or VAR, Model 2: Intercept (no trend) in CE, no intercept or trend in VAR, Model 3: Intercept in CE and VAR, no trends in CE and VAR, Model 4: Intercept in CE and VAR, linear trend in CE, no trend in VAR, Model 5: Intercept and quadratic trend in the CE intercept and linear trend in VAR.

<sup>13</sup> See Asteriou and Hall (2007).

<sup>14</sup> For more detail see Asteriou and Hall (2007).

## CHAPTER V

### EMPIRICAL ANALYSIS

#### 5.1 Data

The quarterly data covering the period of 1980: 01 – 2007: 04 for 14 industrialized countries is used for the empirical analysis. The countries consist of Australia, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Spain, Sweden, Switzerland, United Kingdom and United States. The composition of the sample is restricted due to unavailability of data and the necessity for a balanced panel. The variables include money supply, income, interest rates, price levels, and nominal exchange rates.

As a measure of money supply, M3 expressed in millions of national currency is used for all countries except Switzerland and United Kingdom. Money supply is money plus quasi money for Switzerland and United Kingdom due to unavailability of M3 data for these countries.<sup>15</sup> Seasonally adjusted

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<sup>15</sup> See Mark and Sul (2001), Engel and West (2005) and Engel, Mark and West (2007) for the usage of different measures of the money supply across countries. As in Mark and Sul (2001), the money supply data are seasonally adjusted by taking the average of the current and three previous quarters.

industrial production index (2000=100) is utilized as a proxy of income. Three-month deposit rates constitute a measure for short-term interest rates. Price levels are measured using consumer price index (CPI) with base year of 2000. Finally, nominal exchange rates expressed in units of home currency per dollar for 13 industrialized countries are included to the model.

M3, money plus quasi money, industrial production index, and consumer price index data for all countries except Belgium, Italy, Netherlands and Spain are taken from the web site of IMF's International Financial Statistics (IFS). M3 data for Belgium, Italy, Netherlands and Spain and three-month deposit rates data for all countries are taken from Global Financial Database. Nominal exchange rate data expressed in units of home currency per dollar for 13 industrialized countries are taken from the web site of Organization of Economic Co-operation and Development (OECD). With the exception of interest rates all variables are expressed in logarithms.

## **5.2. Panel Unit Root Tests**

This thesis utilizes IPS and Fisher ADF Chi-square panel unit root statistics in order to examine integration properties of the exchange rates, relative income levels, relative money supplies, relative interest rates, and relative price levels across countries. Since LLC and homogeneous type of Hadri test statistics are restrictive in the sense that all cross-sections have or do not have a unit root, focusing on less restrictive IPS and Fisher ADF Chi-square test statistics can lead to more accurate results about integration properties of the variables. IPS

and Fisher ADF Chi-square test results in levels and first differences are presented in Table 5.2.1 and Table 5.2.2, respectively.

**Table 5.2.1: IPS Panel Unit Root Tests for Stationarity**

Variables	Without Trend	With Trend	Conclusion
s	-1.92**	0.02	I(1)
m-m*	-0.89	-1.03	I(1)
y-y*	1.17	1.88	I(1)
i-i*	-2.76***	-0.89	I(1)
p-p*	1.11	-1.15	I(1)
$\Delta s$	-24.35***	-26.15***	I(0)
$\Delta(m-m^*)$	-3.17***	-2.27**	I(0)
$\Delta(y-y^*)$	-13.66***	-14.96***	I(0)
$\Delta(i-i^*)$	-36.26***	-37.03***	I(0)
$\Delta(p-p^*)$	-7.39***	-6.98***	I(0)

**Notes:** The symbols s, m-m\*, y-y\*, i-i\*, and p-p\* denote nominal exchange rate, money supply differences, industrial production index differences, deposit rates differences, and consumer price index differences of individual countries and US, respectively. Similarly,  $\Delta s$ ,  $\Delta(m-m^*)$ ,  $\Delta(y-y^*)$ ,  $\Delta(i-i^*)$ , and  $\Delta(p-p^*)$  stand for first differences of nominal exchange rate, money supply differences, industrial production index differences, deposit rates differences, and consumer price index differences of individual countries and US, respectively. All variables are expressed in logarithms except interest rates. IPS test statistics includes an individual intercept and both an individual trend and intercept. Lag lengths are chosen by Akaike Information Criterion (AIC). \*\*, and \*\*\* stand for the level of significance at 5 %, and 1%, respectively.

**Table 5.2.2: ADF-Fisher Chi-square Panel Unit Root Tests for Stationarity**

Variables	Without Trend	With Trend	Conclusion
s	39.90**	25.22	I(1)
m-m*	32.11	32.68	I(1)
y-y*	23.24	14.02	I(1)
i-i*	45.61**	29.41	I(1)
p-p*	31.58	42.67**	I(1)
$\Delta s$	515.02***	502.21***	I(0)
$\Delta(m-m^*)$	47.72***	38.86**	I(0)
$\Delta(y-y^*)$	272.94***	285.16***	I(0)
$\Delta(i-i^*)$	714.22***	652.74***	I(0)
$\Delta(p-p^*)$	131.52***	114.84***	I(0)

**Notes:** The symbols s, m-m\*, y-y\*, i-i\*, and p-p\* denote nominal exchange rate, money supply differences, industrial production index differences, deposit rates differences, and consumer price index differences of individual countries and US, respectively. Similarly,  $\Delta s$ ,  $\Delta(m-m^*)$ ,  $\Delta(y-y^*)$ ,  $\Delta(i-i^*)$ , and  $\Delta(p-p^*)$  stand for first differences of nominal exchange rate, money supply differences, industrial production index differences, deposit rates differences, and consumer price index differences of individual countries and US, respectively. All variables are expressed in logarithms except interest rates. ADF Fisher Chi-square test statistics includes an individual intercept and both an individual trend and intercept. Lag lengths are chosen by Akaike Information Criterion (AIC). \*\*, and \*\*\* stand for the level of significance at 5 %, and 1%, respectively.

According to IPS test results, all variables with individual intercepts and trends are nonstationary in levels at 1 % and 5 % significance levels. IPS test results also indicate that all variables with individual intercepts are nonstationary in

levels at 1 % significance level except interest rate differences. According to ADF Fisher Chi-square test results, all variables with and without trends are nonstationary at 1 % significance level. Both IPS and Fisher Chi-square test statistics indicate that all variables with and without trends in first differences are stationary at 5 % significance level. Overall, all variables are found to be integrated of order one,  $I(1)$ , meaning that a prerequisite for the panel cointegration analysis is provided. Based on panel unit root statistics results one can proceed to panel cointegration analysis in order to examine long-run relationship between exchange rates and monetary fundamentals or exchange rates and Taylor rule fundamentals.

### **5.3. Panel Cointegration Analysis**

This thesis employs Johansen Fisher panel cointegration test in order to provide evidence for the existence of a long run relationship between exchange rates and monetary fundamentals or Taylor-rule fundamentals across countries. Johansen Fisher panel cointegration test is applied for the panel version of (9) and (16). Pedroni and Kao panel cointegration tests are based on Engle-Granger (1987) methodology which is quite restrictive when analyzing the cointegrating properties of an  $n$ -dimensional vector of  $I(1)$  variables where several cointegration relationships may arise. Since this thesis estimates a trivariate system including nominal exchange rates, price levels and interest rates, it is possible to obtain more than one cointegrating relationship formed by these variables. Johansen Fisher Panel Cointegration technique has an advantage over

the Pedroni and Kao panel cointegration tests in the sense that it relaxes the assumption of a unique cointegrating vector among the variables.

Before applying Johansen Fisher Panel Cointegration test, an optimal lag length for two different panel-based VAR models should be determined; the first one consists of nominal exchange rates, money supply differences and income level differences and the second one includes nominal exchange rates, interest rates differences and price level differences. After choosing optimal lag length based on Schwarz information criteria, the Pantula principle is used for both models in order to determine the appropriate model regarding the deterministic components. Table 5.3.1 and Table 5.3.2 show Pantula principle results for monetary model including nominal exchange rates, relative money supplies and relative income levels and Taylor-rule model including nominal exchange rates, relative interest rates and relative price levels, respectively.

**Table 5.3.1: The Pantula Principle Results using  $s$ ,  $(m-m^*)$ ,  $(y-y^*)$**

r	n-r	Model 2	Model 3	Model 4
0	3	74.48	75.98	89.67
1	2	37.53**	37.47	44.16
2	1	24.29	34.96	22.81

$s$  : Nominal exchange rate expressed in units of home currency per dollar  
 $m-m^*$  : Difference of money supply between individual countries and US  
 $y-y^*$  : Difference of industrial production index between individual countries and US  
 $r$  : Number of cointegrating relations

$n$  : Number of variables entering the Vector Autoregression (VAR)  
 Model 2: Intercept (no trend) in CE, no intercept or trend in VAR  
 Model 3: Intercept in CE and VAR, no trends in CE and VAR  
 Model 4: Intercept in CE and VAR, linear trend in CE, no trend in VAR  
 \*\* indicates the first time that the null can not be rejected at 5 % significance level according to probabilities computed using Chi-square distribution.  
 Lag length is chosen as six based on Schwarz information criteria.

**Table 5.3.2: The Pantula Principle Results using  $s$ ,  $(i-i^*)$ ,  $(p-p^*)$**

$r$	$n-r$	Model 2	Model 3	Model 4
0	3	117.9	123.2	138.7
1	2	45.81	60.36	33.08**
2	1	26.32	67.08	21.13

$s$  : Nominal exchange rate expressed in units of home currency per dollar  
 $i-i^*$  : Difference of deposit rates between individual countries and US  
 $p-p^*$  : Difference of consumer price index between individual countries and US  
 $r$  : Number of cointegrating relations  
 $n$  : Number of variables entering the Vector Autoregression (VAR)  
 Model 2 : Intercept (no trend) in CE, no intercept or trend in VAR  
 Model 3 : Intercept in CE and VAR, no trends in CE and VAR  
 Model 4 : Intercept in CE and VAR, linear trend in CE, no trend in VAR  
 \*\* indicates the first time that the null can not be rejected at 5 % significance level according to probabilities computed using Chi-square distribution.  
 Lag length is chosen as six based on Schwarz information criteria.

Starting with the smaller number of cointegrating vectors  $r = 0$ , the model selection procedure is based on checking whether the trace statistic for model 2 rejects the null, if yes proceeding to the right, checking whether third model rejects the null, and so on. The results in Table 5.3.1 indicate that the appropriate model for the panel cointegration analysis of the exchange rate with

monetary fundamentals is the one with the intercept (no trend) in CE, no intercept or trend in VAR, namely model 2.

The results in Table 5.3.2 also show the appropriate model for the panel cointegration analysis of the exchange rate with Taylor-rule fundamentals is the one with intercept in CE and VAR, linear trend in CE, no trend in VAR, namely model 4. Hence, based on these findings, Johansen Fisher panel cointegration results of the exchange rate with monetary fundamentals and the exchange rate with Taylor-rule fundamentals are given in Table 5.3.3 and Table 5.3.4, respectively.

**Table 5.3.3: Johansen Fisher Panel Cointegration Results using  $s$ , ( $m-m^*$ ), ( $y-y^*$ )**

Number of Cointegrating Vectors	Fisher Stat. from Trace Test		Fisher Stat. from Max- Eigen Test	
	Prob.	Prob.	Prob.	Prob.
$r \leq 2$	24.29	0.55	24.29	0.55
$r \leq 1$	37.53	0.06	32.8	0.16
$r \leq 0$	74.48**	0.00	60.1**	0.00

$s$  : Nominal exchange rate expressed in units of home currency per dollar  
 $m-m^*$  : Difference of money supply between individual countries and US  
 $y-y^*$  : Difference of industrial production index between individual countries and US

$r$  : Number of cointegrating vectors

\*\* denotes statistical significance at the 1 % level.

EViews6 computes probabilities using asymptotic Chi-square distribution.

Lag length is equal to six.

**Table 5.3.4: Johansen Fisher Panel Cointegration Results using  $s$ ,  $(i-i^*)$ ,  $(p-p^*)$**

Number of Cointegrating Vectors	Fisher Stat. from Trace Test	Prob.	Fisher Stat. from Max-Eigen Test	Prob.
$r \leq 2$	21.13	0.74	21.13	0.74
$r \leq 1$	33.08	0.16	28.08	0.35
$r \leq 0$	138.70**	0.00	138.00**	0.00

$s$  : Nominal exchange rate expressed in units of home currency per dollar

$i-i^*$  : Difference of deposit rates between individual countries and US

$p-p^*$  : Difference of consumer price index between individual countries and US

$r$  : Number of cointegrating vectors

\*\* denotes statistical significance at the 1 % level.

EViews6 computes probabilities using asymptotic Chi-square distribution.

Lag length is equal to six.

Table 5.3.3 indicates that there exist a unique cointegrating vector among nominal exchange rates, relative money supplies, and relative income levels at 1 % significance level. Similarly, Table 5.3.4 shows that existence of one cointegrating vector among nominal exchange rate, relative interest rates, and relative price levels at 1 % significance level. The conclusion that there is one cointegrating vector, however, does not necessarily imply support for the monetary model or Taylor rule model in determining the exchange rates. For that to be present, the relationship among the variables must be reasonably consistent with those implied by the monetary and the Taylor rule exchange rate equations.

On the basis of these results, two long-run exchange rate equations can be estimated in order to assess the impact of monetary and Taylor-rule fundamentals on nominal exchange rates.

#### **5.4. Estimation and Results**

In the cointegrated panels, using ordinary least square (OLS) method to estimate the long-run equation leads to biased and inconsistent estimator of the parameters. OLS estimates suffer from asymptotic bias unless the regressors are strictly exogenous, so that the OLS standard errors can not generally be used for valid inference. Pedroni (2000) proposes fully modified ordinary least square (FMOLS) estimation while Kao and Chiang (2000) and Mark and Sul (2001) recommend the dynamic ordinary least squares (DOLS) as the alternative methods of panel cointegration estimation.

FMOLS estimation corrects for endogeneity and serial correlation to the ordinary least square (OLS) estimator. To correct for endogeneity bias and to obtain an unbiased estimator of the long-run parameters, DOLS uses a parametric adjustment to the errors by augmenting the static regression with leads, lags, and contemporaneous values of the regressors in first differences. Both FMOLS and DOLS provide consistent estimates of standard errors that can be used for inference.

According to Kao and Chiang (2000) FMOLS and DOLS estimators have normal limiting properties, even though DOLS estimator outperforms FMOLS estimator in empirical analysis. On the basis of the earlier findings in

favor of panel DOLS estimation, DOLS method is employed to estimate long-run exchange rate equation which relates nominal exchange rates with monetary and Taylor-rule fundamentals. Then, the following exchange rate equation based on panel DOLS method is estimated:

$$s_{it} = \alpha_i + \theta_t + x_{it}'\beta + \sum_{j=-q}^q c_{ij}\Delta x_{it+j} + v_{it} \quad (23)$$

where  $i = 1, \dots, N$  and  $t = 1, \dots, T$ ,  $\{s_{it}\}$  are 1x1,  $\beta$  is a 2x1 vector of slope parameters,  $\{\alpha_i\}$  are the intercepts indicating individual fixed effects,  $\{\theta_t\}$  stand for common time effects,  $\{v_{it}\}$  are the error terms,  $\{x_{it}\}$  are 2x1 where  $x_{1it} = (m_{it} - m_t^*)$  and  $x_{2it} = (y_{it} - y_t^*)$  for the monetary model,  $x_{1it} = (i_{it} - i_t^*)$  and  $x_{2it} = (p_{it} - p_t^*)$  for the Taylor-rule model and  $q$  stands for number of leads and lags of the first differenced regressors.

Table 5.4.1 and Table 5.4.2 denote panel DOLS results for monetary model and Taylor-rule model, respectively.

**Table 5.4.1: Panel Dynamic OLS Estimates of Monetary Model**

	FE	FE-T
$\hat{\beta}_1$	0.118*** (0.023)	0.290*** (0.018)
$\hat{\beta}_2$	0.332*** (0.046)	0.163*** (0.063)

$\hat{\beta}_1$  : Coefficient of money supply differences between individual countries and US

$\hat{\beta}_2$  : Coefficient of industrial production index differences between individual countries and US

\*\*\* indicates significance at 1 % level.

Standard errors for the coefficient estimates are given in parenthesis.

FE: Fixed effect in the regression

FE-T: Fixed effect together with a time dummy in the regression.

Based on Schwarz information criterion, FE model is estimated with four lags and three leads, FE-T model is estimated with four lags and two leads.

**Table 5.4.2: Panel Dynamic OLS Estimates of Taylor-rule Model**

	FE	FE-T
$\hat{\alpha}_1$	-0.023*** (0.001)	-0.011*** (0.001)
$\hat{\alpha}_2$	0.985*** (0.043)	1.007*** (0.03)

$\hat{\alpha}_1$  : Coefficient of deposit rate differences between individual countries and US

$\hat{\alpha}_2$  : Coefficient of consumer price index differences between individual countries and US

\*\*\* indicates significance at 1 % level

Standard errors for the coefficient estimates are given in parenthesis.

FE: Fixed effect in the regression

FE-T: Fixed effect together with a time dummy in the regression

Based on Schwarz information criterion, FE model is estimated with four lags and two leads, FE-T model is estimated with four lags and three leads.

According to monetary model,  $\beta_1$  should be equal to 1 and  $\beta_2$  should be less than zero. However, the coefficient estimates in Table 5.4.1 are not in line with

the theory even though they are statistically significant at 1 % level. When money supply of home country increases relative to that of foreign country by one percentage, home currency depreciates by 0.11 percentages. Similarly, one percent increase in income level of home country relative to that of foreign country leads to depreciation of home currency by 0.33 percentages. When income of home country increases relative to that of foreign country, home country's consumption might increase. If the consumption of the home country mostly depends on the goods imported from the foreign country, rises in imports result in the depreciation of the home currency.

On the other hand, the Taylor-rule model implies that both  $\alpha_1$  and  $\alpha_2$  are between zero and one. Table 5.4.2 indicates that coefficient estimate for price level differences between home and foreign countries in the model with fixed effects has the correct sign implied by the theory and it is statistically significant at 1 % level. However, the coefficient estimate for interest rate differences between home and foreign countries does not have the correct sign although it is statistically different from zero at 1 % level. The coefficient estimate for the price level indicates that when the price level of home country increases relative to foreign country, home currency depreciates.

The negative coefficient estimate for the interest rate indicates that when the interest rate of home country increases relative to foreign country by one unit, the log of exchange rate decreases by 0.023 units implying the appreciation of the home currency. This result can be interpreted as follows: Increases in the interest rate of home currency relative to foreign currency lead to increase the

demand for the domestic assets which in turn cause home currency to appreciate relative to foreign currency.

Overall, there is no support for the monetary model and there is little support for Taylor-rule model in explaining exchange rates. Even though both monetary and Taylor-rule fundamentals determine exchange rates in the long-run, the effects of these fundamentals on exchange rates are not the same as what the theory implies.

## **CHAPTER VI**

### **CONCLUSION**

In this thesis, Johansen Fisher panel cointegration technique is employed to test the validity of the monetary and the Taylor-rule models of exchange rate for US dollar exchange rates over 1980:01-2007:04 periods for 13 industrialized countries. Unlike the findings of Engel and West (2005) which is a study of time series, this thesis finds a cointegration relationship between the nominal exchange rates and a set of fundamentals implied by the monetary model and the Taylor rule model. The panel cointegration results regarding the existence of a long-run relationship between the nominal exchange rates and the monetary fundamentals are in line with the findings of Groen (2000), Mark and Sul (2001), Rapach and Wohar (2004), and Crespo-Cuaresma et al. (2005).

Based on the panel cointegration test results, the cointegrating coefficient estimates for the monetary and the Taylor rule models are found by using panel dynamic ordinary least square (DOLS) estimation. The results for the monetary model indicate that the coefficient of the money supply has the correct sign but

its magnitude is not close to one, the coefficient of the income level has the wrong sign even though it is statistically significant.

The results for the Taylor-rule model show that the coefficient of the price level has the correct sign and size and the coefficient of the interest rate has the wrong sign although it is statistically significant.

The earlier studies which provide the evidence for a cointegrating relationship between monetary fundamentals and the nominal exchange rates also find cointegrating coefficients which agree with the values suggested by the monetary model. This thesis finds different results from the previous studies because it uses a longer span of the data compared to previous studies. Moreover, it estimates the cointegrating vectors with a different estimation technique which has not been used previously to examine the behavior the US dollar exchange rate for the industrialized countries.

Groen (2000) and Rapach and Wohar (2004) provide noteworthy results regarding the cointegrating coefficients of the monetary model for the industrialized countries. Groen (2000) covers the period of 1973:1-1994:1-4 and he uses least square dummy variable (LSDV) to estimate cointegrating coefficients. Similarly, Rapach and Wohar (2004) utilizes the period of 1973:1-1997:1 and they estimate the cointegrating coefficients by using more than one estimation technique including LSDV, pooled mean group estimator (PMGE) and dynamic SUR estimates. On the other hand, Mark and Sul (2001) cover the period of 1973:1-1997:1 and they assume pre-specified values for the cointegrating coefficients of  $\beta_1 = 1$  and  $\beta_2 = -1$  before using panel

cointegration tests. However, the estimation results of this thesis provide evidence for the invalidity of the assumption of Mark and Sul (2001).

This thesis follows Crespo-Cuaresma et al. (2005) in estimating panel cointegrating coefficients of the monetary model and Taylor rule model. Using the Euro exchange rate across Central and Eastern European countries, Crespo-Cuaresma et al. (2005) find support for the cointegrating coefficients of the monetary model by using DOLS estimator. However, this thesis shows that the same estimation technique does not provide evidence for the validity of the monetary model if one uses US dollar exchange rate for the industrialized countries. Overall, the findings of this thesis imply that there is no support for the monetary model and there is little support for Taylor-rule model in explaining exchange rates.

As a future study, the other estimation techniques such as fully-modified OLS and pool mean group estimation (PMGE) can be used to estimate two long-run nominal exchange rate equations in order to see whether the cointegrating coefficient estimates change in favor of the monetary model and the Taylor-rule model.

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## APPENDIX

Derivation of equation (7)

$$s_t = (1 + b_2)^{-1} x_t + b_2(1 + b_2)^{-1} E_t(s_{t+1})$$

$$s_{t+1} = (1 + b_2)^{-1} x_{t+1} + b_2(1 + b_2)^{-1} E_{t+1}(s_{t+2})$$

$$E_t(s_{t+1}) = (1 + b_2)^{-1} E_t(x_{t+1}) + b_2(1 + b_2)^{-1} E_t(s_{t+2}), \quad \text{using the fact that}$$

$$E_t(E_{t+1}(s_{t+2})) = E_t(s_{t+2})$$

and plugging this to the first equation:

$$s_t = (1 + b_2)^{-1} x_t + b_2(1 + b_2)^{-2} E_t(x_{t+1}) + b_2^2(1 + b_2)^{-2} E_t(s_{t+2}), \quad \text{applying the same procedure iteratively to } E_t(s_{t+2}), E_t(s_{t+3}) \text{ and imposing transversality condition,}$$

$$\lim_{i \rightarrow \infty} [b_2 / (1 + b_2)]^i E_t(s_{t+i}) = 0, \quad \text{one can obtain}$$

$$s_t = (1 + b_2)^{-1} \sum_{i=0}^{\infty} [b_2 / (1 + b_2)]^i E_t(x_{t+i})$$

Derivation of equation (8)

Firstly rewrite  $x_t$  as:

$$x_t = \frac{1}{1 + b_2} \sum_{i=0}^{\infty} \left[ \frac{b_2}{1 + b_2} \right]^i x_t$$

Then subtract  $x_t$  from the both sides of equation (7) to obtain

$$s_t - x_t = \frac{1}{1+b_2} \sum_{i=1}^{\infty} \left[ \frac{b_2}{1+b_2} \right]^i E_t(x_{t+i} - x_t)$$

$$s_t - x_t = \frac{1}{1+b_2} \sum_{i=1}^{\infty} \left[ \frac{b_2}{1+b_2} \right]^i \sum_{j=1}^i E_t(\Delta x_{t+j})$$

$$s_t - x_t = \frac{1}{1+b_2} \sum_{j=1}^{\infty} E_t(\Delta x_{t+j}) \sum_{i=j}^{\infty} \left[ \frac{b_2}{1+b_2} \right]^i$$

$$s_t - x_t = \sum_{j=1}^{\infty} \left[ \frac{b_2}{1+b_2} \right]^j E_t(\Delta x_{t+j})$$