Effects of Fundamentals on the Exchange Rate: A Panel Analysis for a Sample of Industrialised and Emerging Economies

José Eduardo de A. Ferreira*
Central Bank of Brazil

Abstract: This paper tests the traditional monetary model of exchange rates for a sample of industrialized and emerging market economies by making use of panel techniques that allow for a high degree of heterogeneity across countries. The results demonstrated partial support for the monetary model for industrialised market economies but not for emerging ones. This constitutes a puzzle as it would expect countries with greater monetary instability to show a stronger association between exchange rates and monetary fundamentals.

JEL Classification: **F31, F37, F41.**

Keywords: Foreign Exchange, Fundamentals, Panel Data, Unit Roots, Assets.

Acknowledgements: I am grateful to Dr. Miguel Leo-Ledesma and Prof. Alan Carruth, for helpful suggestions on earlier drafts. I also owe thanks to Dr. Mathan Satchi, Prof. Michael Bleaney and Dr. Matloob Piracha for final comments.

* **Address for Correspondence:** Department of International Reserves, Central Bank of Brazil, Brasilia -Brazil, Tel. +55 61 3414 1846 (email: jeafdf@yahoo.com.br).

Effects of Fundamentals on the Exchange Rate: A Panel Analysis for a Sample of Industrialised and Emerging Economies

Introduction

The literature on exchange rate economics over the past three decades has witnessed many controversial debates with respect to the most appropriate model to be used in empirical studies. A considerable number of studies in recent years has developed more sophisticated econometric methods, but the results still continue to be at best tentative in explaining exchange rate movements. In fact, many empirical and theoretical studies have been unable to give a convincing explanation of exchange rate movements. Consequently, exchange rate economics still remains a challenging field for researchers [see Flood and Rose (1999) and Rogoff, (1999)].

Perhaps one of the most important areas of study on exchange rates is the ability of economic fundamentals to explain exchange rate behaviour. Despite considerable research input and a plethora of empirical results for a range of countries, the results leave a number of issues unresolved.

Some empirical work has focused on the analysis of the cointegrating properties between nominal exchange rates and monetary fundamentals using panel data techniques [see, for example, MacDonald et al. (2003), Rapach and Wohar (2004), Mark and Sul (2001), Groen (2000)]. The results suggest some support for the view that economic fundamentals have long-run effects on the exchange rate. Husted and MacDonald (1998) also find evidence in favour of the monetary model of the exchange rate using multi-country panel data, allowing for a limited amount of cross-country heterogeneity. Cushman (2000), on the other hand, using the Johansen (1991) approach, concludes that there is no support for the monetary model in US-Canadian data as the cointegrating coefficients differ significantly from those predicted by the theory.

Rogoff (1999) states that inflation rates in industrialised economies such as the United States, Germany and Japan have tended to converge downwards towards zero. It is more difficult to identify the effects of monetary shocks on exchange rates. In contrast, emerging economies are historically more likely to present weak economic fundamentals such as ever-expanding budget deficits, high inflation rates, low rates of economic growth, excess supply of money etc. As such, the mismanagement of economic aggregates is believed to generate direct effects on exchange rate behaviour [see, for example, Rogoff (1996), Bahmani-Oskooee and Kara (2000), Moosa (2000) and Civcir (2002)]. Such adverse economic circumstances are believed to be a key ingredient in favour of the monetary approach to exchange rates[1].

Table 1 displays some comparative economic figures between emerging and industrialised countries. The figures for the inflation rate and budget balance suggest that emerging countries are more subject to monetary shocks which should impact on exchange rates.

Table 1 – Some Economic Indicators for Industrialised and Emerging Countries 1975-2000*

Economic Indicators**	Industrialised Countries	Emerging Countries
Domestic Debt (%GDP)	34.1	45.7
Budget Balance (%GDP)	-2.9	-4.6
Economic Growth	2.7	3.4
Inflation Rate (CPI)	6.8	34.1

Source: World Bank Database.
* The figures belong to countries used in this study.
** Period Averages.

The second contribution relies on the use of panel unit roots and panel cointegration analyses applied distinctively to the two groups of economies. For panel unit root detection, in particular, two types of tests are used: Levin, Lin, and Chu - LLC (2002) and Im, Pesaran and Shin - IPS (2003). The basic difference between the LLC test and the IPS test is that the former is characterised by constraining the coefficient of the lagged variable in the ADF regression to be homogeneous across all units. It means that the only relevant source of heterogeneity is the unit-specific fixed effect. The IPS test is an extension of the LLC test allowing for heterogeneity of the coefficients on the lagged variables and using the group-mean *t*-student statistics to test the null hypothesis. The tests for panel cointegration, in turn, are based on Pedroni (1995, 1997, 1999) and assumes the null hypothesis of no cointegration. The main advantage of this method is that it allows for considerable heterogeneity in the panel. The heterogeneity in this approach includes the possibility of assuming heterogeneous slope coefficients, fixed effects and individual specific deterministic trends. All testing procedures are applied distinctively to the two sets of economies.

[1] Bleaney *et al.* (1999) show that unit root models are more appropriate to model mean reversion in real exchange rates for high inflation countries.

The third contribution is associated with the estimation technique, which is based on important findings published initially by Robertson and Symons (1992), later by Pesaran and Smith (1995) and deepened by Im, Pesaran and Smith (1996). These authors, basically, showed the inconsistency of estimates produced by pooled or aggregated data when the context is strongly heterogeneous. As this inconsistency can lead to spurious regressions, the use of group mean estimators is recommended. This procedure aims, fundamentally, at eliminating the bias caused by the covariance between the error term and regressors by averaging across groups. The pooled mean group (PMG) estimator developed by Pesaran, Shin and Smith (1999), adopted in this paper, offers this facility and may overcome problems of inconsistent estimates. This estimator offers considerable flexibility being appropriate to a heterogeneous panel environment and can model nominal exchange rate movements successfully. The objective is essentially to generate unbiased and consistent estimates. Although PMG assumes the long-run coefficients to be the same, its statistical properties allow the intercept, short-run coefficients and error variances to differ freely across units. Furthermore, this econometric method also provides an additional advantage in comparison with the mean group (MG) estimation by allowing that certain parameters are the same across units [see Pesaran and Smith (1995) for more details].

The econometric estimation procedure adopted in this paper is also supported by results provided in Phillips and Moon (1999) and Kao (1999). These results reveal that some panel datasets made of reasonable N (number of groups) and T (length of the sample period) can produce consistent estimates of long-run average parameters even if panel time series are not cointegrated [error process I(1)] Roughly, Phillips and Moon (1999) state that the estimation of a long-run relation among variables is feasible even for cases where the use of the pure time series dimension alone suggests that the regression may be spurious. The underlying intuition is that the introduction of a cross-section dimension, as well as the averaging across groups, reduces the noise in the relationship, that is, the potential covariance between the error term and regressors. Thus, even if no cointegration vectors are found due to, for instance, the presence of bubbles, these estimators will still provide consistent estimates of long-run elasticities.

Four recent studies analyse the monetary approach to the nominal exchange rate using panel data techniques. The first one, developed by Groen (2000), applies OLS and assumes homogeneity for the money supply and output coefficients. Groen uses dummy variables to capture level changes and employs the Levin and Lin (1993) panel unit root test that implies homogeneity in the ADF regression parameters. The second study of Mark and Sul (2001) uses a panel dynamic OLS estimator in which, by construction, has individual-specific or time-specific effects, but slope coefficients are homogeneous across groups. Rapach and Wohar (2004) make an extensive investigation of the long-run monetary model of exchange rate determination on a country-by-country and panel data basis for a large number of industrialised countries. They only find clear evidence in favour of the monetary model using panel procedures, and assuming homogeneity of cointegrating coefficients. Finally, MacDonald et al. (2003) develop a panel data study for six selected Central and Eastern countries with heterogeneous exchange rate regimes and show that the monetary model of exchange rates provides an acceptable explanation of exchange rate behaviour.

While using the standard monetary model of exchange rates, the investigation strategy of this paper differs from these cited studies by allowing for either homogeneity or heterogeneity for panel unit root and panel cointegration tests across units. Heterogeneity is assumed for the short-run coefficients and homogeneity for the long-run coefficients across units. Furthermore, the sample of countries sample in this study is based on homogeneous exchange rate regimes according to the criterion provided by Reinhart and Rogoff (2002). The analysis for industrialised market economies is based on freely floating exchange rate regimes, whereas for emerging market economies is based on managed exchange rate regimes.

In summary, this paper is structured into five distinct sections. Section 2 outlines the main features of the theoretical model underlying the empirical analysis. Section 3 describes the statistical data, variables used and sources. Section 4 presents a theoretical overview of the panel unit root test, panel cointegration test and the estimation method as well as the empirical results. Finally, Section 5 presents the conclusions.

2 - Theoretical Model

According to the economic literature, the flexible-price monetary approach, as originally developed by Frenkel (1976), Mussa (1976), Johnson (1977) and Bilson (1978), assumes the quantity theory of money and PPP condition as the building blocks for developing economic models. Although the flexible-price monetary model was the dominant model in that period, it rapidly gained unpopularity due to its poor empirical performance in explaining exchange rate determination. This weakness led to the development of the sticky-price or overshooting exchange rate model of Dornbusch

(1976). Nevertheless, this is a controversial issue as there is a significant number of studies published on monetary models of exchange rate determination which make use of either flexible-price or sticky-price assumptions (see Sarno and Taylor, 2002). This paper uses three different monetary model versions in which the exchange rate determinants are analysed in flexible-price format following Frankel (1979)[2].

The strategy of this paper is to begin from a simple flexible-price monetary model and assume a symmetric treatment between domestic and foreign countries[3]. As usual, the exchange rate is defined as the relative price of two monies. A brief description of the flexible-price monetary model in discrete time is presented as follows[4]:

$$m_t = p_t + \kappa y_t - \theta i_t \tag{1}$$

$$m_t^* = p_t^* + \kappa^* y_t^* - \theta^* i_t^* \tag{2}$$

equations (1) and (2) represent the monetary equilibrium for both domestic and foreign countries, in which m_t denotes the money supply, p_t the price level, y_t the income level and i the interest rate.

As the PPP condition is assumed to hold continuously (flexible price context), it implies that movements in the exchange rate must be directly proportional to movements in prices in both the short run and the long run, that is:

$$s_t = p_t - p_t^* \tag{3}$$

Given the monetary equilibrium expressed by (1), (2) and the PPP condition denoted by (3), and if the money supply determines the price level, according to the quantity theory of money, then indirectly the nominal exchange rate is also determined by the relative money supply, or the money supply differential, represented by:

$$s_t = \left(m_t - m_t^*\right) - \left(\kappa y_t - \kappa^* y_t^*\right) + \left(\theta i_t - \theta^* i_t^*\right) + e \tag{4}$$

---

[2] The introduction of a sticky-price framework within the flexible-price monetary model aims, on the one hand, at taking into account deviations from the PPP condition, and the other hand inserting a mechanism which reflects the short-run behaviour of the exchange rate determined by expectations of the long-run inflation rate differential. This procedure is not adopted in this paper.

[3] Note that similar theoretical frameworks were used more reecently by Cushman (2000), Mark and Sul (2001), Groen (2000), Groen (2002), Rapach and Wohar (2004) to cite a few. This approach can also be found in Obstfeld and Rogoff (1996), Hallwood and MacDonald (2000) and Sarno and Taylor (2002).

[4] Note that all variables are in the logarithm form except the interest rate.

Equation in (4) is called the fundamental equation of the exchange rate for the flexible-price monetary model. Assuming, for simplicity, that the income elasticities and interest rate semi-elasticities of money demand do not differ significantly between countries ($\kappa = \kappa^* \; and \; \theta = \theta^*$), then equation (4) takes the following form:

$$s_t = \left(m_t - m_t^*\right) - \kappa\left(y_t - y_t^*\right) + \theta\left(i_t - i_t^*\right) + e_t \qquad (5)$$

Equation (5) denotes the long-run equation for exchange rate determination, and the effects of m, y and i can be explained by taking into account the economic agents' expected behaviour. Note that the uncovered interest rate parity - UIRP condition $(\Delta^e s_{t+1} = i_t - i_t^*)$ does not hold in equation (5)[5]. Thus, an increase in the domestic money stock leads market agents to spend further on goods and services and drives prices up. Assuming that the PPP condition always holds, the exchange rate is expected to depreciate as a consequence of higher money stock and price level. This depreciation is proportional and means that the coefficient on $\left(m_t - m_t^*\right)$ equals one if the monetary model holds. In contrast, given the domestic money stock, a higher real domestic income level implies a corresponding increase in money demand, and a decrease in the price level is required to maintain the monetary equilibrium. In this case there will be an appreciation of the exchange rate.

The sign on the interest rate differential coefficient in (5) reflects the price regime in which the country is operating. A statistically significant positive coefficient on the interest rate differential reveals that the price regime is based on flexible prices. It supports a flexible-price context as the exchange rate is affected positively by domestic price level changes. On the other hand, if a statistically significant negative coefficient on the interest rate differential is found, then the conclusion favours a sticky-price context. The negativity of this coefficient implies that the exchange rate is determined by monetary policy changes and the sticky price regime holds (see Frankel, 1979). Basically, the idea is that a positive sign on the interest rate differential coefficient implies the exchange rate depreciates as a result of rises in the domestic price level. When the domestic price level increases the real money supply decreases and the home nominal interest rate increases relative to the interest rate abroad. If the exchange rate

[5] The uncovered interest rate parity (UIRP) condition may not hold in presence of market frictions and extraneous noises not associated to exchange rate fundamentals.

depreciates at the same proportion to the rise in prices the flexible price regime holds (Chicago Theory). On the other hand, a negative sign on the interest rate differential coefficient implies the exchange rate appreciates as result of a tight monetary policy and a sticky price regime takes place. The rise in the home interest rate relative to the abroad one due a tight monetary policy leads to an increase in capital inflow and a consequent exchange rate appreciation at least in the short run. If prices are sticky in the short run, then a tight monetary policy does not lead an instantaneous fall in the price level and the real money supply remains unchanged. The increase in the interest rate generates an exchange rate appreciation (Keynesian approach).

A restricted version of the monetary approach can still be derived if the UIRP condition holds. Thus, the UIRP condition can be invoked from equation (5), that is:

$$\Delta^e s_{t+1} = \left(i_t - i_t^*\right) \tag{6}$$

where the expected change in the exchange rate ($\Delta^e s$) is proportional to the nominal interest rate differential. Equation (5) can be suitably rearranged as follows:

$$s_t = \left(1+\theta\right)^{-1}\left(m_t - m_t^*\right) - \kappa\left(1+\theta\right)^{-1}\left(y_t - y_t^*\right) + \theta\left(1+\theta\right)^{-1} s_{t+1}^e + e_t \tag{7}$$

Applying the rational expectations solution to (7) generates the following representation:

$$s_t = \left(1+\theta\right)^{-1}\sum_{i=0}^{\infty}\left(\frac{\theta}{1+\theta}\right)^i E_t\left[\left(m_{t+i} - m_{t+i}^*\right) - \kappa\left(y_{t+i} - y_{t+i}^*\right)\right] + \varepsilon_t \tag{8}$$

where $\varepsilon_t = \phi\left[E_t\left(s_{t+i} - s_t\right)\right] + e_t$. If no bubbles are present in the stochastic process then $\varepsilon_t \to I\left(0\right)$. This means that there is a fundamental solution. Equation (8) constitutes a second version of equation (5) assuming that the UIRP condition holds.

Finally, from equation (5) and following Frankel (1979), it is also possible to derive a third version based on the hypothesis of PPP condition holds in the long run. The third version considers the use of inflation expectations by market agents to denote a second fundamental assumption. It assumes that expected changes in the exchange rate are a function of the gap between the current exchange rate and its long-run equilibrium value plus the expectations about domestic and abroad inflation rates. In practice, the version consists of introducing the expected inflation rate differential

$\left(\pi - \pi^*\right)$ in (5) and examining the statistical significance of its parameter. The objective of introducing the expected inflation rate is to test if the model is a long-run model or a short-run model. This procedure can initially be derived by taking the expression as follows:

$$\Delta^e s = -\varphi\left(s - \overline{s}\right) + \pi - \pi^* \tag{9}$$

where π and π^* are the current expectations of long-run inflation rates at home and abroad, respectively. Equation (9) means that expected changes in the exchange rate must be proportional to the current gap in the short run $\left(s - \overline{s}\right)$, and once $s = \overline{s}$ in the long run, the expected change must be proportional to the expected long-run inflation rate differential $\left(\pi - \pi^*\right)$.

Substituting the UIRP condition into equation (9) and rearranging, the following result is obtained:

$$s - \overline{s} = -\varphi^{-1}[\left(i - \pi\right) - (i^* - \pi^*)] \tag{10}$$

Equation (10) defines the exchange rate gap as a function of the real interest rate differential. Thus, in the long-run, when $s_t = \overline{s}_t$, then $\overline{i} - \overline{i}^* = \pi - \pi^*$, where $\overline{i}$ and $\overline{i}^*$ denote the long-run interest rates. Given that the purchasing power parity (PPP) condition holds continuously and $(\overline{i} - \overline{i}^*) = (\pi - \pi^*)$, the expression for the long-run exchange rate (5) can alternatively be defined as follows:

$$\overline{s} = \overline{p} - \overline{p}^*$$

$$\overline{s} = \left(\overline{m} - \overline{m}^*\right) - \kappa\left(\overline{y} - \overline{y}^*\right) + \theta\left(\pi - \pi^*\right) + e \tag{11}$$

If it is also assumed that the equilibrium money supply and income levels are defined by their current levels, equation (11) can be introduced into (10) to obtain the following final equation:

$$s = \left(m - m^*\right) - \kappa\left(y - y^*\right) - \varphi^{-1}\left(i - i^*\right) + \left(\varphi^{-1} + \theta\right)\left(\pi - \pi^*\right) + e \tag{12}$$

From equation (12) two conclusions can be drawn. Firstly, φ^{-1} is expected to be significantly negative for a sticky-price regime to hold and $\left(\varphi^{-1} + \theta\right)$ will equal zero.

Secondly, φ^{-1} is expected to be significantly positive for a flexible-price regime to hold and $\left(\varphi^{-1}+\theta\right)$ will equal zero[6].

The equations set out in (5), (8) and (12) comprise the economic models investigated in this paper:

Model I : $ner = \delta_1 (m - m^*) - \delta_2 (y - y^*) + e$

Model II : $ner = \delta_1 (m - m_t) - \delta_2 (y - y^*) + \delta_3 (ir - ir^*) + e$

Model III : $ner = \delta_1 (m - m^*) - \delta_2 (y - y^*) + \delta_3 (ir - ir^*) + \delta_4 (\pi - \pi^*) + e$

with $\delta_1 = 1, \delta_2 = \kappa < 0$, $\delta_3 = \varphi^{-1} > 0$ for the flexible-price regime and $\delta_3 < 0$ for the stick-price regime, and $\delta_4 = \left(\varphi^{-1}+\theta\right) = 0$.

3 - Data

The data are collected from the International Financial Statistics provided by the International Monetary Fund with end-of-quarter periodicity. They consist of two data sets: a ten-countries set for industrialised market economies and a seven-countries set for emerging market economies. The basic criterion to select countries was based on average inflation rate and exchange rate regimes taking into account data availability and large exchange rate bands for managed regimes. For the group of industrialised market economies the period extends from 1980:1 to 1998:4 as from 1999:1 onwards a single currency system (Euro) was in place. For the emerging market economies the data extends from 1992:1 to 2002:2 when the non-fixed exchange rate series for most countries are available. Particularly, for emerging market economies, the sample period is associated with the need to conduct the analysis for a non-fixed exchange rate regime. The exchange rate regime follows the classification provided by Reinhart and Rogoff (2002). The selected emerging market economies, except Chile, adopt the exchange rate regimes based on crawling peg regimes with large bands of exchange rate changes (on average 10% bands).

[6] Note that $(\varphi^{-1}+\gamma)$ is expected to equal zero as the exchange rate should be entirely determined by monetary shocks in the long run. The statistical non-significance of this coefficient implies market participants have already incorporated the long run inflation rate into their expectations, and thus the equation 3.1.5 is empirically confirmed as the long-run equation for the exchange rate determination.

The set of industrialised market economies is composed of Australia, Canada, France, Germany, Italy, Japan, Norway, Portugal, Spain, and the United Kingdom. The emerging market economies comprise Chile, South Korea, Malaysia, Mexico, South Africa, Thailand and Turkey.

This study conducts the analysis using the nominal exchange rate (ER) against the US dollar country by country as the endogenous variable. The regressors denoting money supply consist of the logarithm of the broad money supply differential ($m2 - m2^{*} = M2$) and the narrow money supply differential $M1$ constructed in a similar way. The use of a broad money supply $M2$ allows the examination of its impact on the exchange rate based on a concept of lower liquidity assets. The other regressors are the logarithm of the real income differential ($gdp - gdp^{*} = Y$) denoted by GDP in volume, the interest rate differential ($ir - ir^{*} = IR$) based on the money market rate and the expected long-run inflation rate differential (π) formed by the consumer price index-CPI ($\pi - \pi^{*} = \prod$).

The variable denoting the expected long-run inflation rate differential is constructed by averaging the quarterly consumer price index over the preceding year, that is, a moving average proxy (see Frankel, 1979). Finally, the differentials are calculated by assuming the United States as numeraire country denoted by an asterisk[7].

4 – Econometric Methodology

The procedure of empirical investigation tests the theoretical models as in Section 2 by developing the following basic steps:

- Analysis of unit roots: The panel tests are based on Levin et al. (2002) and Im et al. (2003) which allow for panel homogeneity and heterogeneity across units, respectively, to the coefficient on the lagged variable;

- Analysis of cointegration: The panel tests for cointegration follow Pedroni (1999) which allows for panel heterogeneity and different cointegranting vectors across units;

- Estimation of cointegration vectors: The econometric estimation procedure assumes the results found by Phillips and Moon (1999) and uses the pooled

---

7 Note that all variables are in logarithm form except the interest rate.

mean group (PMG) estimation developed by Pesaran et al. (1999) to estimate the cointegration vectors.

This procedure of investigation is equally applied to the samples of industrialised economy countries and emerging economy countries as described in Section 2.

4.1- Analysis of Unit Roots in Panel Data

According to Levin, Lin and Chu (2002), inferences about the existence of unit roots, and cointegration as well, can be made more powerful by including a cross-section dimension. The addition of a cross-section dimension can work, in special conditions, as repeated draws from the same distribution. Researchers have also observed that time and cross-section dimensions together increase the power of the panel test statistics and allow distributions of estimators to converge to normality [see Baltagi and Kao (2000) and Banerjee (1999)].

Levin, Lin and Chu (2002) formulated a panel-based unit root test procedure which can still incorporate individual specific intercepts and time trends. This procedure allows the residual variance and the pattern of higher-order serial correlation to vary freely across units. The test is designed to evaluate the null hypothesis that each unit in the panel has integrated residuals against the alternative hypothesis that all units have stationary residuals. The underlying intuition is that as both cross-section and time series dimensions are enlarged, the regression estimators as well as the test statistics embody both asymptotic properties of stationary panel data and the asymptotic properties of integrated time series data. An interesting feature is that, as opposed to non-normal distributions of unit root test statistics for single time series, the panel unit root test statistic has a limiting normal distribution.

Thus, for modelling purposes, it is assumed that the stochastic process is composed of a panel of units $i = 1,\ldots,N$ each of them is observed over time periods $t = 1,...,T$. The final objective is to determine whether this process is integrated, that is, contains a unit root for each group in the panel. Similarly to the case of a single time series (see Dickey and Fuller,1981), an intercept and time trend for each unit can be included. Additionally, it is assumed that all units in the panel have identical first-order partial correlation, but it is also permitted for all parameters of the disturbance process to vary freely across units. Finally, the rationale of the test consists of analysing, under the null hypothesis, whether each individual time series has a unit root and, under the alternative hypothesis, the process $\{y_{it}\}$ is trend-stationary for each unit in the panel.

Nevertheless, the Levin, Lin and Chu's specification restricts the coefficient on the lagged level variable in the unit root test to be homogeneous across units. Im, Pesaran and Shin (2003), thus, suggested allowing for heterogeneity related to that coefficient under the alternative hypothesis. This suitable modification aims at capturing the individual realities in panel studies based on a large number of units[8].

The results of both Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003) tests are reported in Tables 2 and 3 for variables in levels and for variables in first difference, respectively.

Table 2 – Panel Unit Root Tests: Levels [1]

Variables (In level)		Industrialised Economies		Emerging Economies	
		LLC test	IPS Test	LLC test	IPS Test
ER	No trend	-1.70639*	-3.49072**	2.3702	2.4002
	Trend	-1.70523*	-2.45601**	-0.22211	-0.07283
M2	No trend	0.72587	0.47114	0.98656	-0.16911
	Trend	-0.84174	-1.39407	-0.31366	-0.20887
M1	No trend	1.40149	-0.06773	-0.95107	-2.13951
	Trend	-1.29783	-1.69963	-0.44270	-0.75707
Y	No trend	1.2494	0.06889	-0.12171	-1.91276*
	Trend	-0.80508	0.91644	0.06755	-1.13305
IR	No trend	-1.69997*	-2.4424**	-1.90889*	-3.78502**
	Trend	-1.17553	-1.29126	-2.42927**	-5.29192**
Π	No Trend	-0.29289	-1.96485**	2.15877	2.7237
	Trend	-0.24864	-0.11621	-0.43637	0.44206

[1]Critical Values: * = - 1.645 (5% level) , ** = -1.947 (1% level).

[8] The LLC and IPS test procedures were applied by O'Connell (1998) to explain the failure of purchasing power parity-PPP to hold in the presence of market frictions. Also Papell (1997) using data for 21 industrialised countries from 1973 to 1994 did not reject the unit root null at 10% level for PPP. MacDonald (1996), Coakley and Fuertes (1997), O'Connell (1998), Papel and Theodoridis (1998), Ru-Lin Chiu (2002) also used these test procedures just to cite a few. The Levin, Lin and Chu test has also been employed by Frankel and Rose (1996), Oh (1996) and Lothian (1996). They tested the PPP hypothesis using panel data and some of them found evidence supporting this hypothesis.

Table 3 – Panel Unit Root Tests: 1st Difference [1]

Variables (In 1st difference)		Industrialised Economies		Emerging Economies	
		LLC test	IPS Test	LLC test	IPS Test
ER	No trend	Stationary	Stationary	-7.83054**	-10.12823**
	Trend	Stationary	Stationary	-6.27176**	-9.54109**
M2	No trend	-7.01133**	-9.93503**	-8.32932**	-10.50587**
	Trend	-5.86114**	-9.17421**	-6.65449**	-10.80993**
M1	No trend	-10.1154**	-10.76055**	-7.14308**	-8.94996**
	Trend	-9.36826**	-11.99829**	-5.48417**	-8.00195**
Y	No trend	-11.614**	-15.30639**	-6.25675**	-8.46943**
	Trend	-8.97043**	-14.09973**	-4.77875**	-7.61463**
IR	No trend	Stationary	Stationary	Stationary	Stationary
	Trend	-8.85097**	-12.20578**	Stationary	Stationary
Π	No Trend	-1.71090*	-2.35205**	-1.78308*	-1.96739**
	Trend	-3.13952**	-3.39179**	-1.67325*	-1.87065*

[1]Critical Values: * = - 1.645 (5% level) , ** = -1.947 (1% level).

In general, most of the results reported in Table 2 confirmed the expectations. The variables in levels were found mostly to be non-stationary at the 5% and 1% levels. These results are valid either for panels with trends included or not.

The exchange rate for industrialised market economies, unexpectedly, was found to be stationary in levels for both types of tests. Some hypotheses arise to explain this unexpected result: (1) the power of the cross-section dimension prevailed in order to attenuate the non-stationary effects implicit in many time-series; (2) the use of a fixed lag order in the ADF regression; (3) the possibility of cross section dependence effects take place (see Pesaran, 2003). For emerging market economies the expectation is confirmed so that the exchange rate is found to be non-stationary for both versions of the panel unit root test[9].

The money supply, based both on the broad concept M2 and the narrow concept M1, is found to be non-stationary in levels at the 5% and 1% significance levels, even if a heterogeneous trend is included for industrialised and emerging market economies. The objective was to examine if the exclusion of lower liquidity assets, included in the M2 concept, might bring significant changes to the results. The interest rate differential series were found to be stationary for industrialised economies, except when a heterogeneous trend is included, and stationary for emerging economies. This finding may reveal that the uncovered interest rate parity (UIRP) condition holds, and thus models II and III (see section 2) do not hold for both sets of countries. Finally, for the

[9] Note, nonetheless, that it is used a fixed lag order p_i=4 in the ADF regression. The tests were also applied by using lag order p = 1, 2, 3, but no significant difference was detected.

expected long-run inflation rate series the LLC and IPS tests indicate non-stationary behaviour in levels for both types of economies.

The series in first difference were all found to be stationary with only one suitable differentiation. Table 3 reports the results obtained from the LLC and IPS tests at the 5% and 1% significance levels[10].

The unexpected stationarity for the nominal exchange rate for industrialised market economies suggested that the use of a fixed lag order $p_i = 4$ in the ADF regression might lead to misleading conclusions[11]. Again the IPS test is applied to the exchange rate for industrialised market economies, but by allowing different lag orders for each country. The traditional procedure to identify the most appropriate lag order consists of running successively the ADF regression country by country and obtaining the maximum statistically significant lag order p_i[12]. Next, the standard IPS formula is employed for different values of $E\left\{t_{iT}\left(p_i\right) | \delta_i = 0\right\}$ and $Var\left\{t_{iT}\left(p_i\right) | \delta_i = 0\right\}$ tabulated by Im *et al*. (2003) for different *Ts* and lag orders. Table 4 displays the lag order used country to country.

Table 4 – Maximum Lag Order in the ADF Regression

Countries	Exchange Rate	
	No Intercept	Intercept
Australia	3	3
Canada	3	3
France	1	4
Germany	1	1
Italy	1	1
Japan	1	1
Norway	3	3
Portugal	1	3
Spain	3	3
United Kingdom	7	7

[10] Note also that the inclusion of a trend did not change this conclusion.
[11] The advisable rule is to set a lag order so that it should be relatively small in order to save degrees of freedom, but large enough to eliminate a possible autocorrelation in the error process.
[12] The lag order p_i can be different across groups, so the method proposed by Hall (1990) for selecting the appropriate lag order can be used. Basically, it consists of , for a given sample length T, choosing a maximum lag order using *t*-statistic on $\hat{\theta}_{iL}$ to determine if a smaller lag order can be more suitable.

The procedure generated the statistics $\Psi_{\bar{t}} =$ -2.61586 and $\Psi_{\bar{t}} =$-2.37492 without and with heterogeneous intercept included, respectively. The results confirmed the stationarity in the nominal exchange rate movements for industrialised market economies either at the 5% or 1% levels of significance[13]. Nevertheless, this stationarity for nominal exchange rates may be a misleading result as heterogeneous panels are supposed to present cross section dependence which affects standard panel unit root tests. To deal with this difficulty, Pesaran (2003) proposes a simple testing procedure that may asymptotically eliminate the effects of cross sectional dependence in the series. Basically, his approach consists of augmenting standard DF and ADF unit root tests with cross section averages of lagged levels and first differences of the individual series as the following representation:

$$\Delta y_{it} = \alpha_i + b_i y_{i,t-1} + c_i \bar{y}_{t-1} + \sum_{j=0}^{p} d_{ij} \Delta \bar{y}_{t-j} + \sum_{j=1}^{p} \Delta y_{i,t-j} + \varepsilon_{it} \qquad (13)$$

where $\bar{y}_{t-1}$ and $\Delta \bar{y}_{t-j}$ in (13) are cross section averages of lagged levels and first differences, respectively. In essence, by averaging the *t*-ratios for the coefficient b_i for individual series, denoted by individual CADF_i statistics, generates a modified version of the *t*-bar test proposed by Im, Pesaran, Shin ($CIPS = N^{-1} \sum_{i=1}^{N} CADF_i$) for panel data analysis.

The CIPS test for the exchange rate considering the lag orders displayed in Table 4 generated CIPS_1 = -1.788 and CIPS_2 = -1.764 without and with intercept included, respectively. Theses results imply the nominal exchange rate for industrialised market economies is, in fact, non-stationary at 1% level of significance and the previous results of stationarity may be affected by cross section dependence[14]. The non-stationarity of the series may lead to a long-run relationship between the exchange rate and its fundamentals as predicted by the theory.

---

[13] One explanation for this unexpected result may rely on the fact that there may be a very high correlation of the nominal exchange rate across countries. When cross-sectional correlation takes place the properties of panel tests are violated so that misleading conclusion may emerge.

[14] Pesaran (2003) shows that the limit distribution of the CIPS statistic exists and is free of nuisance parameters. The critical values at 1% level of significance to the present analysis are CIPS_1= -1.95 and CIPS_2= -2.54 without and with intercept included, respectively.

4.2 - Analysis of Cointegration in Panel Data

Additional developments within the empirical and theoretical literature in econometrics have led to the development of new methods for testing for cointegration in panel data. The Pedroni's approach (see Pedroni, 1995, 1997) assumes the null hypothesis of no cointegration, and it is linked to the ideas of the pioneering work of Engle and Granger (1987)[15]. Pedroni (1997), taking into account the concern for working with stationary time series in levels, developed several tests for the null hypothesis of no cointegration in panel data. The main contribution of his approach is to allow for considerable heterogeneity. Basically, Pedroni (1997) constructed asymptotic distributions for test statistics based on heterogeneous dynamics across units, endogenous regressors, fixed effects and individual-specific deterministic trends for bivariate regressions. Furthermore, Pedroni (1997) reinforces this heterogeneous panel approach by including appropriate tests either for common autoregressive roots or for heterogeneity across units in autoregressive roots under the alternative hypothesis.

The Pedroni (1997) method is somewhat limited by focusing only on simple bivariate regressions. Thus, Pedroni (1999) extended it and developed a test for the null hypothesis of no cointegration for the case with multiple regressors. In fact, the null hypothesis of the Pedroni (1999) test statistic is that the variables are not cointegrated for each unit in the panel. The alternative hypothesis is that there exists only one cointegrating vector for each unit in the panel, although it may differ across units.

For the analysis proposed in this paper, four statistics out of seven developed by Pedroni are selected and reported. The first one is non-parametric and it is analogous to the Phillips and Perron (1988) *t*-statistic, the second is a parametric statistic and it is analogous to the familiar augmented Dickey-Fuller *t*-statistic. The other two statistics are based on a group mean approach and they are also analogous to the Phillips and Perron *t*-statistic and augmented Dickey-Fuller *t*-statistic. The selection criterion followed comparative advantage analyses of each statistic based on the underlying data-generating process (see Pedroni, 1999 for more details)[16].

The statistical properties demonstrated that the test statistics converge to the standard normal form $N(0,1)$ under the null hypothesis and diverges to negative infinity

[15] The general idea of the Engle-Granger method consists of estimating the long-run relationship between two (or more) variables, and then inserting the deviations from the long-run path, lagged appropriately, as the error correction mechanism into the short-run equation.

[16] Note that, all statistics are constructed using residuals generated by the ADF regression and by the use of nuisance parameters estimators.

under the alternative hypothesis. In effect, Pedroni (1995, 1997) shows that when both the T and N dimensions grow large the individual member statistics of the panel cointegration converge to normal distributions by virtue of conditional independence across the i members. In practice, the statistics' interpretation leads to the rejection of the null hypothesis of no cointegration whenever large negative values are found.

The panel cointegration tests are performed on the three different theoretical models described in Section 2. The test statistics consider the possibility of two alternative statistical approaches. The first one, the panel-test statistic, or the within-dimension statistic, assumes a common coefficient γ_i in the error process for both Phillips-Perron and ADF regressions across units in the panel. The second one, by contrast, the group-test statistics, or the between-test statistics, assumes the possibility of a different coefficient γ_i across units. Note once again that as the test statistics diverge to negative infinity under the alternative hypothesis, large negative values imply the rejection of the null hypothesis.

The analysis of the results is carried out using the broad and narrow concepts of money supply (M2 and M1, respectively). For industrialised market economies, the results using M2 demonstrate a clear rejection of the null hypothesis of no cointegration at the 5% and 1% significance level, based on the pp-statistics (see Table 5).

Table 5 – Panel Cointegration Test (M2 for Industrialised Economies)

Models		**Statistics**			
		Panel pp-stat	**Panel ADF-stat**	**Group pp-stat**	**Group ADF-stat**
Model I[1]	No trend	-1.7342*	-0.8857	-2.7547**	-1.20834
	Trend	-1.6847*	0.43146	-2.9997**	0.18148
Model II[2]	No trend	-0.50297	-0.10291	-2.6732**	-1.16895
	Trend	-2.5121**	0.76852	-4.4698**	-0.21497
Model III[3]	No trend	-2.6342**	0.21114	-4.1468**	-0.32914
	Trend	-4.9405**	-0.84005	-6.8116**	-2.12615**

Critical Values: * = - 1.645 (5% level) , ** = -1.947 (1% level) .
[1] **Model I : $er = (m2-m2^*) - (y - y^*) + e$**
[2] **Model II : $er = (m2-m2^*) - (y - y^*) + (ir - ir^*) + e$**
[3] **Model III : $er = (m2-m2^*) - (y - y^*) + (ir - ir^*) + (\pi - \pi^*) + e$**

In effect, the panel cointegration tests applied to industrialised market economies generated statistically significant Phillips-Perron statistics for all models, except for model II with no trend, and based on panel statistics (panel pp-stat) and group statistics (group pp-stat). The inclusion of heterogeneous trends did not change the results substantially. This result leads to the conclusion that nominal exchange rate movements for industrialised market economies are cointegrated with monetary fundamentals. In this case, the monetary approach is appropriate to explain the long-run movements in nominal exchange rates. The results revealed that there is a long-run relationship between exchange rate movements and monetary fundamentals. It is still important to note that the larger absolute values for group statistics may lead to an additional conclusion that the cointegrating vectors differ across countries.

Nonetheless, this conclusion is entirely based on the Phillips-Perron tests, through both the panel PP-statistics and the group PP-statistics. The conclusion for cointegration did not hold when the Augmented Dickey-Fuller *t*-statistic tests (ADF-stat) were applied to the sample. The panel-statistics and the group-statistics, using a fixed lag order (k =4)[17] in the ADF regression, produced results statistically insignificant at both levels of significance. It may suggest that a serially correlated error process might be present leading to the rejection of the null hypothesis of no cointegration in the previous analysis based on the PP-statistics. Hence, the conclusion in favour of cointegration between the exchange rate and its fundamentals in industrialised market economies may be misleading.

For emerging market economies, the results displayed in Table 6 demonstrate, unequivocally, the non-existence of cointegration between the exchange rate and monetary fundamentals for all models tested. The tests, either for panel statistics or group statistics, both using the PP-test and ADF-test, revealed that the traditional monetary model of the exchange rate does not hold for emerging market economies, that is, long-run movements in the exchange rate are not associated with movements in monetary fundamentals (see Table 6).

Table 6 – Panel Cointegration Test (M2 for Emerging Economies)

Models		**Statistics**			
		Panel pp-stat	**Panel ADF-stat**	**Group pp-stat**	**Group ADF-stat**
Model I[1]	No trend	0.38004	0.31876	1.94724	2.15738
	Trend	1.65999	1.42158	2.03549	1.67311
Model II[2]	No trend	0.42217	0.64665	1.91327	2.13392
	Trend	-2.1547**	-0.79145	-2.4986**	0.38478
Model III[3]	No trend	0.38861	1.32885	0.5111	0.32414
	Trend	-1.77286*	-0.0158	-0.54592	-0.2304

Critical Values: * = - 1.645 (5% level) , ** = -1.947 (1% level) .
[1] Model I : er = (m2-m2*) – (y – y*) + e
[2] Model II : er = (m2-m2*) – (y – y*) + (ir – ir*) + e
[3] Model III : er = (m2-m2*) – (y – y*) + (ir – ir*) + (π – π *) + e

[17] Similar test procedure was applied for lag order k=3, k=2 and k=1, and the results did not differ significantly.

This is an unexpected result since this class of economies is subject to monetary shocks more frequently with adverse effects on the exchange rate. Theoretically, the monetary approach to exchange rates should be appropriate to explain exchange rate movements in countries with profiles of high inflation rates. Although models II and III demonstrated some evidence of cointegration reported by statistically significant panel PP-statistics, this result may be due to the inclusion of heterogeneous trends which may produce distorted effects on the regression error. The general conclusion is that the null hypothesis of no cointegration cannot be rejected for emerging market economies at the 5% and 1% level of significance.

The same cointegration tests are also applied to the two sets of countries, but now using a narrow concept of money supply M1. Tables 7 and 8 display the results obtained for industrialised economies and emerging economies, respectively[18]. The use of M1 as the money supply measure reinforced the hypothesis of no long-run relationship between the exchange rate and fundamentals so that no evidence of cointegration was found for both sets of industrialised and emerging market economies at both levels of significance (see Tables 7 and 8).

Table 7 – Panel Cointegration Test (M1 for Industrialised Economies)

Models		Statistics			
		Panel pp-stat	Panel ADF-stat	Group pp-stat	Group ADF-stat
Model I[1]	No trend	-0.1264	-0.37002	0.24150	-0.48273
	Trend	-0.13676	1.54556	-0.51128	1.86844
Model II[2]	No trend	0.36883	0.30945	0.59918	0.68433
	Trend	-0.14199	2.32360	-0.79118	2.60654
Model III[3]	No trend	0.81511	2.52654	1.15356	3.33194
	Trend	-0.47995	2.48842	-1.41895	2.84885

Critical Values: * = - 1.645 (5% level) , ** = -1.947 (1% level) .
[1] Model I : er = (m1-m1*) – (y – y*) + e
[2] Model II : er = (m1-m1*) – (y – y*) + (ir – ir*) + e
[3] Model III : er = (m1-m1*) – (y – y*) + (ir – ir*) + (π – π *) + e

[18] Note that, once again, the lag order for the ADF regression is fixed to be k=4 in order to keep a standard strategy of investigation.

Table 8 – Panel Cointegration Test (M1 for Emerging Economies)

Models		Statistics			
		Panel pp-stat	Panel ADF-stat	Group pp-stat	Group ADF-stat
Model I[1]	No trend	-1.83183*	3.37106	-2.12968**	4.34182
	Trend	-1.88045*	-1.04853	-1.03453	-0.59187
Model II[2]	No trend	-0.04141	3.84923	-1.03624	4.70848
	Trend	-1.58517	-0.46867	-0.51591	0.82264
Model III[3]	No trend	0.14393	2.22686	-0.00305	2.18307
	Trend	-3.0118**	-2.51079**	-2.91793**	-1.57446

Critical Values: * = - 1.645 (5% level) , ** = -1.947 (1% level) .
[1] **Model I : er = (m1-m1*) – (y – y*) + e**
[2] **Model II : er = (m1-m1*) – (y – y*) + (ir – ir*) + e**
[3] **Model III : er = (m1-m1*) – (y – y*) + (ir – ir*) + (π – π *) + e**

4.3 - Method of Econometric Estimation

The results of the previous section raise an important issue when the hypothesis of cointegration does not hold, which relates to the third contribution proposed for this paper. Phillips and Moon (1999) demonstrated that long-run average relations between integrated panel vectors can be found even if no cointegration is detected by pooling data. The pooling data technique appears to meet Phillips and Moon (1999)'s points better as data on the different units over several periods of time are gathered within a same model.. As compared to a single cross-section, the pooling data procedure provides an advantage by relaxing assumptions that are implicit in cross-sectional analysis.

An efficient econometric method of estimation designed to overcome the potential bias in panel cointegrated regression models is the pooled mean group (PMG) estimator proposed by Pesaran, Shin and Smith - PSS (1999). Basically, the PMG estimation method consists of an intermediate procedure between the traditional pooled regression method which assumes that the slope coefficients and error variances are identical, and the mean group (MG) estimator which estimates N separate regressions and averages the coefficients. Pesaran and Smith (1995) point out that the major difficulty related to the MG estimator is that it does not take into account that in panel regression models some parameters may be the same across units. Although the MG estimator provides consistent estimates of the mean long-run parameters, they will be inefficient if long-run homogeneity holds. The PMG estimator imposes constraints on the long-run coefficients by assuming them to be identical, but allows for short-run coefficients and error variances to differ freely across units.

The PMG estimator also offers the possibility of setting different dynamic specifications across units, such as the number of lags included in the model for example, without imposing equality of the short-run coefficients. Furthermore, it also assumes that different units are supposed to be influenced by common factors (technologies, budget constraints, arbitrage conditions etc.) set up in a long-run homogeneous framework. Thus, it supposes that equilibrium relationships between variables are similar across different units in the long run. Accordingly, under a long-run slope homogeneity assumption, the PMG estimator is able to provide consistent and efficient estimates.

Hence, if there is actually a long-run relationship between the exchange rate and its monetary fundamentals, it is expected that the use of the PMG estimation procedure is able to produce statistically significant estimates of the long-run coefficients. Section 4.3.1 presents the main empirical findings from applying the PMG procedure to the exchange rate and monetary fundamentals.

4.3.1 – PMGE and Exchange Rate Fundamentals

The pooled mean group (PMG) estimation procedure is applied to the theoretical model discussed in Section 2. Basically, the empirical model for the monetary approach to the exchange rate follows a similar structure to that employed by Pesaran, Shin and Smith (1999), who examined the standard consumption function in Davidson *et al.* (1978) for a sample of OECD countries. The proposal is to analyse three different empirical models, including also an interpretation of the interest rate and the expected future inflation rate differentials in the exchange rate equation. The maximum lag order is 3 which allows enough time for monetary fundamentals disturbances to affect movements in the exchange rate[19]. Finally, an empirical autoregressive distributed lag (ARDL) equation is constructed for the three specifications presented in Section 2. The full long-run exchange rate specification is given by:

$$er_{it} = \theta_{0i} + \theta_{1i}\left(m_{it} - m_{it}^{*}\right) - \theta_{2i}\left(y_{it} - y_{it}^{*}\right) + \theta_{3i}\left(ir_{it} - ir_{it}^{*}\right) + \theta_{4i}\left(\pi_{it} - \pi_{it}^{*}\right) + u_{it} \qquad (14)$$

$$i = 1, 2, \ldots, N \quad and \quad t = 1, 2, \ldots, T$$

[19] Note that the lag orders k=2 and k=1 were also applied and the results did not differ significantly.

where er_{it} is the nominal exchange rate and $\left(m_{it}-m_{it}^{*}\right)$, $\left(y_{it}-y_{it}^{*}\right)$, $\left(ir_{it}-ir_{it}^{*}\right)$ and $\left(\pi_{it}-\pi_{it}^{*}\right)$ correspond, respectively, to the differential between the domestic and foreign money supply, the real income, the interest rate and the expected future inflation rate. According to the theory, the coefficient θ_{1t} is expected to be one if the monetary approach to the exchange rate holds. Once this hypothesis holds, the PMG estimation procedure enables the estimation of a common long-run coefficient and additional tests.

The general representation of an ARDL (3,3,3,3,3) is denoted as follows:

$$er_{it}=\mu_i+\lambda_{1iq}\sum_{q=1}^{3}er_{it-q}+\delta_{2iq}\sum_{q=0}^{3}m_{it-q}-\delta_{3iq}\sum_{q=0}^{3}y_{it-q}+\delta_{4iq}\sum_{q=0}^{3}ir_{it-q}+\delta_{5iq}\sum_{q=0}^{3}\pi_{it-q}+\varepsilon_{it} \tag{15}$$

and the general error correction representation - ECM of (15) is given by:

$$\Delta er_{it}=\phi_i\left(er_{it-1}-\theta_{0i}-\theta_{1i}\,m_{it-1}+\theta_{2i}\,y_{it-1}-\theta_{3i}\,ir_{it-1}-\theta_{4i}\,\pi_{it-1}\right)+\delta_{1iq}\sum_{q=1}^{3}\Delta er_{it-q}+\delta_{2iq}\sum_{q=0}^{3}\Delta m_{it-q}$$
$$-\delta_{3qi}\sum_{q=0}^{3}\Delta y_{it-q}+\delta_{4qi}\sum_{q=0}^{3}\Delta ir_{it-q}+\delta_{5qi}\sum_{q=0}^{3}\Delta\pi_{it-q}+\varepsilon_{it} \tag{16}$$

where from (16):

$$\theta_{0i}=\frac{\mu_i}{1-\lambda_{i1}L^1-\lambda_{i2}L^2-\lambda_{i3}L^3}\quad;\qquad\theta_{1i}=\frac{\delta_{10i}+\delta_{11i}L^1+\delta_{12i}L^2+\delta_{13i}L^3}{1-\lambda_{i1}L^1-\lambda_{i2}L^2-\lambda_{i3}L^3}$$

$$\theta_{2i}=-\frac{\delta_{20i}+\delta_{21i}L^1+\delta_{22i}L^2+\delta_{23i}L^3}{1-\lambda_{i1}L^1-\lambda_{i2}L^2-\lambda_{i3}L^3}\quad;\qquad\theta_{3i}=\frac{\delta_{30i}+\delta_{31i}L^1+\delta_{32i}L^2+\delta_{33i}L^3}{1-\lambda_{i1}L^1-\lambda_{i2}L^2-\lambda_{i3}L^3}$$

$$\theta_{4i}=\frac{\delta_{40i}+\delta_{41i}L^1+\delta_{42i}L^2+\delta_{43i}L^3}{1-\lambda_{i1}L^1-\lambda_{i2}L^2-\lambda_{i3}L^3}\quad;\qquad\phi_i=1-\lambda_{i1}L^1-\lambda_{i2}L^2-\lambda_{i3}L^3$$

From equation (16) three different error correction empirical models can be derived, as defined theoretically in Section 2. The three empirical models examined for industrialised and emerging market economies are as follows:

I-

$$\Delta er_{it} = \phi_i \left(er_{it-1} - \theta_{0i} - \theta_{1i}\, m_{it-1} + \theta_{2i}\, y_{it-1} \right) + \delta_{1qi} \sum_{q=1}^{3} \Delta er_{it-q} + \delta_{2qi} \sum_{q=0}^{3} \Delta m_{it-q} - \delta_{3qi} \sum_{q=0}^{3} \Delta y_{it-q} + \varepsilon_{it}$$

II -

$$\Delta er_{it} = \phi_i \left(er_{it-1} - \theta_{0i} - \theta_{1i}\, m_{it-1} + \theta_{2i}\, y_{it-1} - \theta_{3i}\, ir_{it-1} \right) + \delta_{1qi} \sum_{q=1}^{3} \Delta er_{it-q} + \delta_{2qi} \sum_{q=0}^{3} \Delta m_{it-q} - \delta_{3qi} \sum_{q=0}^{3} \Delta y_{it-q}$$

$$+ \delta_{4qi} \sum_{q=0}^{3} \Delta ir_{it-q} + \varepsilon_{it}$$

III-

$$\Delta er_{it} = \phi_i \left(er_{it-1} - \theta_{0i} - \theta_{1i}\, m_{it-1} + \theta_{2i}\, y_{it-1} - \theta_{3i}\, ir_{it-1} - \theta_{4i}\, \pi_{it-1} \right) + \delta_{1qi} \sum_{q=1}^{3} \Delta er_{it-q} + \delta_{2qi} \sum_{q=0}^{3} \Delta m_{it-q}$$

$$- \delta_{3qi} \sum_{q=0}^{3} \Delta y_{it-q} + \delta_{4qi} \sum_{q=0}^{3} \Delta ir_{it-q} + \delta_{5qi} \sum_{q=0}^{3} \Delta \pi_{it-q} + \varepsilon_{it}$$

The economic literature points out to the existence of a positive relationship between the money supply and the exchange rate so that it is expected that the sign on the money supply coefficient is positive and statistically equal to one. The idea is that an increase in the money supply differential leads to a proportional change in the exchange rate. Although the program and the procedures to run the models were the same, some results showed unexpected negative signs on the money supply coefficient. The strategy to deal with this problem was to identify countries in the sample which might be potentially affecting the results. This procedure produced a substantial improvement in the results by dropping Australia and Mexico from the sample for industrialised market economies and emerging market economies, respectively. Additionally, a common trend was also included for industrialized economies which brought about a better result for the estimated coefficients. However, the introduction of a similar common trend for emerging market economies did not bring about significant improvements. Although these two devices were successful in promoting econometric improvements, the results for M2, the broad money supply measure, did not have a coherent economic interpretation. Hence, the outputs for the models I, II, and III, using M2 as a regressor for both sets of countries were excluded from the analysis since they computed negative signs on the money supply coefficient and positive signs on the real income coefficient.

The standard econometric procedures used to run all the different empirical models also included the use of a common deterministic regressor (an intercept). Moreover, in order to prevent the possibility of common factor effects, since lagged dependent variables are used as regressors, the data were cross-section demeaned. The lag orders adopted for industrialised and emerging market economies were based on the Schwarz Bayesian criterion (SBC) on the unrestricted model subject to a maximum lag order of three. For industrialised market economies, using M1 as a regressor, the SBC revealed that the most common lag order are ADRL (1,1,0), ADRL (1,1,0,0) and ADRL (1,1,0,0,0) for the models I, II and III, respectively. For emerging market economies, using M1 as a regressor, the lag order was found to be dispersed, but an ADRL (1,1,0) and (1,1,0,0) prevailed for the empirical models I and II, respectively. The mean group estimates were used as initial estimates of the long-run parameters[20].

The econometric results were organised in three different groups: the first one reports country by country long-run estimates based on panel OLS (group-specific estimation), the second one reports country by country long-run estimates based on PMG, the third one reports the long-run coefficient estimates and long-run relationships of the model based on MG and PMG procedures for the summation of countries as whole.

In practice, for a long-run relationship to hold it is required that ϕ_i (see equation 16) is statistically negative. Hence, for industrialised market economies, using M1 as a regressor and empirical model I, the results based on country by country panel OLS for seven countries, excluding Canada, Italy and the United Kingdom which had insignificant coefficient ϕ_i, supported a long-run relationship between the money supply, real income and the exchange rate at the 5% level of significance. The results of country by country PMG, revealed that all countries had long-run relationships as ϕ_i was found to be statistically significant (see Table 9). This shows that PMG performances better than OLS. Moreover, the long-run relationship estimated by MG and PMG was found to be statistically significant for the summation of industrialised market economies as whole (see Table 9). The long-run coefficients, estimated by MG and by PMG, were positive for M1 and negative for real income for the summations of

[20] The pooled mean group (PMG) computations were carried out using the Newton-Raphson algorithm in a program written in Gauss language.

countries as whole and this result met the expected signs according to the theoretical economic model I (see Table 10).

Table 9 – Long-run Relationship for Industrialised Economies* (by Country)

Countries	Model I		Model II		Model III	
	ϕ_i^{OLS}	ϕ_i^{PMG}	ϕ_i^{OLS}	ϕ_i^{PMG}	ϕ_i^{OLS}	ϕ_i^{PMG}
Canada	-0.0423 (-0.417)	-0.125 (-3.594)	-0.0365 (-0.331)	-0.131 (-3.405)	0.0838 (-0.699)	-0.131 (-3.405)
France	-0.1233 (-2.418)	-0.047 (-2.757)	-0.1187 (-2.046)	-0.054 (-1.896)	-02255 (-3.402)	-0.053 (-1.873)
Germany	-0.1058 (-2.125)	-0.129 (-3.022)	-0.1088 (-2.11)	-0.124 (-2.986)	-0.1802 (-2.677)	-0.118 (-2.861)
Italy	-0.0494 (-1.129)	-0.102 (-2.672)	-0.179 (-3.514)	-0.138 (-3.249)	-0.2006 (-3.598)	-0.14 (-3.294)
Japan	-0.1395 (-2.997)	-0.145 (-4.204)	-0.1353 (-2.775)	-0.154 (-4.169)	-0.1149 (-1.822)	-0.152 (-4.113)
Norway	-0.4331 (-4.394)	-0.427 (-4.496)	-0.4756 (-5.049)	-0.419 (-4.763)	-0.4694 (-4.873)	-0.411 (-4.674)
Portugal	-0.0609 (-2.884)	-0.035 (-2.827)	-0.0662 (-2.713)	-0.048 (-2.146)	-0.0317 (-0.658)	-0.056 (-2.189)
Spain	-0.1322 (-3.111)	-0.075 (-2.422)	-0.1348 (-3.316)	-0.081 (-2.661)	-0.097 (-2.392)	-0.082 (-2.706)
United Kingdom	-0.3774 (-1.638)	-0.075 (-2.501)	-0.1174 (-1.007)	-0.085 (-2.764)	-0.1168 (-1.391)	-0.085 (-2.793)

*** The statistics in parenthesis are *t*-statistics.**

Table 10 – Alternatives Pooled Estimates Using M1 - Industrialised Economies[1]

Model I											
MGE					PMGE					Statistics	
M1	y	ir	π	ϕ_i	M1	y	ir	π	ϕ_i	LR	H-T*
2.667 (1.47)	-5.666 (-1.224)	-	-	-0.143 (-3.583)	1.015 (4.322)	-0.16 (-1.697)	-	-	-0.129 (-3.298)	37.28 [0.002]	1.85 [0.40]
Model II											
MGE					PMGE					Statistics	
M1	y	ir	π	ϕ_i	M1	y	ir	π	ϕ_i	LR	H-T*
3.264 (1.91)	-2.257 (-2.28)	0.02 (1.91)	-	-0.152 (-3.57)	0.888 (3.98)	-0.419 (-2.75)	0.011 (3.25)	-	-0.137 (-3.67)	57.41 [0.000]	5.3 [0.15]
Model III											
MGE					PMGE					Statistics	
M1	y	ir	π	ϕ_i	M1	y	ir	π	ϕ_i	LR	H-T*
2.229 (1.22)	-2.413 (-1.66)	0.027 (2.77)	-1.529 (1.41)	-0.174 (-4.11)	0.901 (3.9)	-0.462 (-2.39)	0.011 (3.24)	-0.082 (-0.35)	-0136 (-3.71)	86.46 [0.0]	5.84 [0.2]

[1] **Figures in parentheses and brackets are *t*-statistics and p-values, respectively.**

* **Hausman Test Statistics.**

The empirical model II includes the interest rate differential as an additional regressor, and it implies that UIRP does not hold. The results of country by country panel OLS showed that Canada and the United Kingdom do not have a long-run relationship between exchange rate movements and monetary fundamentals. In contrast, the results of country by country PMG revealed that all countries have a long-run relationship. The coefficients ϕ_i were all found to be statistically significant (see Table 9). Significant long-run relationships between fundamentals and exchange rate were also found for the summation of the countries as a whole by MG and PMG procedures (see Table 10). Additionally, a statistically significant positive coefficient was found on the interest rate differential by MG and PMG procedures which can be interpreted as evidence in favour of a flexible-price regime (see Table 10)[21]. The other coefficients had the signs predicted by the theoretical economic model.

Empirical model III, which includes the expected future inflation rate differential as an additional regressor, and assumes that UIRP does not hold, produced poorer results from the country by country panel OLS, revealing the inexistence of a long-run relationship for three countries (Canada, Portugal and the United Kingdom). The country by country PMG, on the other hand, produced estimates for the coefficient ϕ_i statistically significant for all countries investigated (see Table 9). Furthermore, and confirming the results obtained for empirical models I and II, MG and PMG procedures revealed once again long-run relationships for industrialised market economies as a whole (see Table 10). The coefficient on the interest rate differential had a positive sign for the summation of developed economies using MG and PMG, supporting the flexible-price context. The other expected signs on coefficients were only correctly estimated by PMG. The coefficient on the expected future inflation rate differential estimated by MG and PMG for the summation of countries as a whole was found to be statistically insignificant (see Table 10). This is an important finding and it reveals that the expected future inflation rate may have already been incorporated into the long-run fundamental economic model of the exchange rate by market agents and captured by the money supply differential. It also confirms the theoretical prediction that expectations of future inflation are not relevant to explain exchange rate movements (see Section 2).

[21] According to the theoretical literature (see Section 2) a positive sign on the interest rate differential coefficient reveals that changes in the price level are accommodated by adjustments on the exchange rate for a given money supply.

The long-run coefficient on the money supply estimated for the summation of industrialised market economies as a whole provided statistically significant values close to one for all economic models. This important result confirms the values predicted by the theoretical model. The estimates on the M1 coefficient computed by PMG were 1.0156 (*t*-statistic for $\theta_{1i} = 1$ is 0.0652), 0.888 (*t*-statistic for $\theta_{1i} = 1$ is -0.5091), 0.9012 (*t*-statistic for $\theta_{1i} = 1$ is -0.4304) for empirical models I, II, III, respectively, which were not found to be statistically different from one as predicted by the theory (see Table 10).

Another relevant finding was computed by the likelihood ratio (LR) statistic and the Hausman test statistic. The LR statistic, which is distributed as a $\chi^2(16)$, $\chi^2(24)$ and $\chi^2(32)$ for empirical models I, II and III, respectively, tests for equal long-run parameters using M1 as a regressor. This statistic rejected the assumption of equal long-run parameters supporting the hypothesis that long-run coefficients differ across countries. The Hausman statistic also tested the same hypothesis based on $V(MGE) - V(PMGE) = 0$ for the three empirical models having M1 as a regressor. The hypothesis of a significant difference for the long-run parameters across countries did not hold (see Table 10). The interpretation given by the Hausman test did suggest that the long-run coefficients do not differ across countries and that there is long-run homogeneity[22].

In relation to emerging market economies, empirical models I and II, using M1 as a regressor, were considered theoretically relevant given the signs computed on the money supply coefficient. Although the results of empirical model III were reported the coefficient signs were not theoretically consistent (see Table 11). Note also that Mexico's data were dropped out from the sample. As opposed to industrialised market economies, the inclusion of a common trend did not produce substantial improvements to the results. Hence, the estimates generated from country by country panel OLS for empirical model I demonstrated that only two countries out of six had a long-run relationship between the exchange rate and monetary fundamentals. Chile and South Africa had coefficients ϕ_i statistically significant at the 5% level of significance for empirical model I with M1 as a regressor. When the data were analysed for the country

[22] The literature usually points out the Hausman test as more robust than LR statistic test.

by country PMG procedure, the results did not show substantial improvements in comparison with panel OLS computations, as just three out of six countries (Chile, South Korea and Turkey) had a statistically significant long-run relationship (see Table 11). For empirical model II, South Korea and South Africa had significant long-run relationships estimated by country by country panel OLS. Chile, Malaysia and Turkey had significant long-run relationships estimated by country by country PMG. For empirical model III only Malaysia did not have a significant long-run relationship for both estimation procedures (see Table 11). Chile was the only country that had a long-run relationship for both estimation procedures, that is panel OLS and PMG, for the three empirical models. The other countries had mixed results which do not allow a robust conclusion. In contrast to the individual computations country by country, significant long-run relationships between exchange rate movements and monetary fundamentals were revealed for the panel of countries as whole for empirical models I and III. PMG computed the coefficient ϕ_i as statistically significant for empirical models I and III and the MG method for empirical model III, exceptionally. Empirical model II did not produce a statistically significant long-run relationship for the summation of emerging market economies as whole by both MG and PMG (see Table 12). The coefficients on the money supply differential were statistically insignificant for empirical models I and II. This was an unexpected result and it may be interpreted that monetary shocks do not explain exchange rate movements in the long run for emerging market economies. Furthermore, empirical model III did not produce the expected sign on the money supply coefficient for both estimation procedures (see Table 12). The coefficients on the real income differential revealed significant negative signs predicted by the theory for both the MG and PMG procedures. The significant positive signs found for the interest rate differential confirmed, once again, that the hypothesis of a flexible-price context also holds for emerging market economies.

Table 11 – Long-run Relationship for Emerging Economies*
(by Country)

Countries	Model I		Model II		Model III	
	ϕ_i^{OLS}	ϕ_i^{PMG}	ϕ_i^{OLS}	ϕ_i^{PMG}	ϕ_i^{OLS}	ϕ_i^{PMG}
Chile	-0.0645 (-1.891)	-0.0423 (-2.286)	-0.0464 (-1.389)	-0.0155 (-1.875)	-0.4003 (-2.164)	-0.3452 (-3.384)
S. Korea	-0.0743 (-1.609)	-0.0735 (-1.751)	-0.0766 (-1.679)	-0.0143 (-1.323)	-0.7582 (-7.619)	-0.687 (-7.596)
Malaysia	-0.0352 (-1.126)	-0.0309 (-1.053)	-0.0253 (-0.911)	-0.027 (-2.183)	-0.204 (-1.568)	-0.1132 (-1.581)
S. Africa	-0.3846 (-2.757)	-0.008 (-1.029)	-0.3851 (-2.704)	-0.0035 (-0.843)	-0.8019 (-5.764)	-0.1806 (-1.968)
Thailand	-0.0163 (-0.676)	-0.0175 (-0.961)	-0.0102 (-0.423)	-0.0143 (-1.598)	-0.9026 (-5.873)	-0.4604 (-4.768)
Turkey	-0.0143 (-1.379)	-0.0149 (-1.954)	-0.0121 (-1.329)	-0.0161 (-2.908)	-0.5045 (-3.714)	-0.9026 (-5.873)

*** The statistics in parenthesis are *t*-statistics.**

Table 12 – Alternatives Pooled Estimates Using M1 - Emerging Economies[1]

Model I											
MGE					PMGE					Statistics	
M1	y	ir	π	ϕ_i	M1	y	ir	π	ϕ_i	LR	H-T*
0.324 (0.56)	-2.637 (-2.543)	-	-	-0.098 (-1.69)	0.002 (0.05)	-2.541 (-4.489)	-	-	-0.031 (-3.17)	11.615 [0.3116]	0.84 [0.66]
Model II											
MGE					PMGE					Statistics	
M1	y	ir	π	ϕ_i	M1	y	ir	π	ϕ_i	LR	H-T*
1.033 (1.14)	-3.371 (-2.64)	0.068 (2.73)	-	-0.019 (-1.37)	1.751 (1.67)	-3.299 (-2.92)	0.095 (2.37)	-	0.003 (0.688)	15.144 [0.4411]	n.a
Model III											
MGE					PMGE					Statistics	
M1	y	ir	π	ϕ_i	M1	y	ir	π	ϕ_i	LR	H-T*
-0.428 (-5.1)	-0.203 (-1.2)	0.004 (1.123)	0.821 (7.08)	-0.595 (-5.41)	-0.424 (-11.7)	-0.371 (-6.6)	0.003 (3.24)	0.898 (7.47)	-0.381 (-4.38)	83.62 [0.00]	n.a.

[1] Figures in parentheses and brackets are *t*-statistics and p-values, respectively.

* Hausman Test Statistics.

It is also of interest to highlight that the likelihood ratio (LR) statistic, distributed as a $\chi^2(10)$, did not reject the hypothesis of equal long-run parameters for empirical models I and II using M1. This result was confirmed by the Hausman test statistic (see Table 12). In essence, it means that the exchange rate movements in emerging market economies are affected similarly in the long run.

Finally, the computations for R^2 and $\overline{R}^2$ of the ECM did not demonstrate a good performance. The three empirical models generated very low values for these two statistics for both sets of economies in the sample. This may lead to the conclusion that the monetary approach to exchange rate determination is not able to capture the degree of variation that exchange rate data shows. Tables 13 and 14 present these computations for industrialised and emerging market economies, respectively[23].

Table 13 – R^2 and $\bar{R}^2$ Estimated by PMG Using M1 - Industrialised Economies

Countries	Model I		Model II		Model III	
	R^2	$\bar{R}^2$	R^2	$\bar{R}^2$	R^2	$\bar{R}^2$
Canada	0.39	0.335	0.38	0.315	0.381	0.304
France	0.373	0.327	0.377	0.322	0.377	0.312
Germany	0.346	0.288	0.344	0.276	0.339	0.259
Italy	0.257	0.148	0.344	0.276	0.258	0.168
Japan	0.541	0.491	0.538	0.481	0.538	0.472
Norway	0.215	0.17	0.308	0.247	0.307	0.235
Portugal	0.222	0.177	0.235	0.179	0.319	0.235
Spain	0.147	0.084	0.118	0.054	0.121	0.043
UK	0.448	0.398	0.457	0.40	0.458	0.392

Table 14 – R^2 and $\bar{R}^2$ Estimated by PMG Using M1 - Emerging Economies

Countries	Model I		Model II		Model III	
	R^2	$\bar{R}^2$	R^2	$\bar{R}^2$	R^2	$\bar{R}^2$
Chile	0.576	0.467	0.497	0.406	n.a.	n.a.
Korea	0.48	0.406	0.485	0.428	n.a.	n.a.
Malaysia	0.529	0.446	0.337	0.263	n.a.	n.a.
S. Africa	0.525	0.417	0.529	0.44	n.a.	n.a.
Thailand	0.428	0.344	0.564	0.50	n.a.	n.a.
Turkey	0.654	0.593	0.564	0.50	n.a.	n.a.

[23] The investigation developed in this paper differs from Rapach and Wohar (2004) by firstly conducting the study for two different categories of economies: industrialised and emerging economies. Secondly, three versions of the monetary model are tested in which two of them assume that UIRP does not hold. This hypothesis is important by assuming that factors others than monetary ones may affect the performance of the model. Thirdly, the long-run inflation rate differential as a third variable tests the validation of the long-run approach. Finally, in contrast to Rapach and Wohar different exchange rate regimes are considered in economic context distinct.

5 - Conclusion

The central aim of this paper was to develop an additional contribution to improve the understanding of the relationship between monetary fundamentals and exchange rate movements. The research strategy consisted of using the standard theoretical monetary model based on traditional variables believed to explain exchange rate behaviour and test it for two different economic environments in panel data format. It was expected that the monetary approach to the exchange rate would have different explanatory power for these distinct economies.

The panel cointegration tests for industrialised market economies showed some evidence, but not robust, of cointegration. The results for emerging market economies rejected the hypothesis of cointegration between exchange rate movements and monetary fundamentals. Hence, the hypothesis that monetary fundamentals determine exchange rate movements was not supported robustly by the cointegration method. The econometric estimation procedures based on the pooled mean group (PMG) for dynamic heterogeneous panels offered more attractive results. The general results for industrialised market economies supported the existence of a long-run relationship between the exchange rate and monetary fundamentals and the coefficients predicted by the theory were found. For emerging market economies, nevertheless, the monetary approach did not provide robust evidence for long-run exchange rate movements. One possible reason for this poorer performance might be associated with the use of managed exchange rate regimes. The adoption of managed exchange rate regimes restricts the performance of the monetary approach in providing an appropriate explanation for exchange rate movements. However, estimated lower values for the coefficient of determination than the theory predicted suggests that the monetary approach is only able to give a partial explanation to exchange rate movements and more robust results are still needed.

In summary, the discussion about the relation between exchange rate behaviour and monetary fundamentals remains open. Even though the results were somewhat mixed, this paper was able to find some support for the monetary model of exchange rate determination. The debate about exchange rate determination has much left to explore. The economic literature points to other possible factors that may be able to explain exchange rate behaviour, such as the presence of bubbles, news, transaction costs, implied risks and other market frictions.

References

Bahmani-Oskooee, M. and Kara, O., 2000,"Exchange Rate Overshooting in Turkey", *Economic Letters 68*, 89-93.

Baltagi, B.H. and Kao, C., 2000, Non-stationary Panels Cointegration in Panel and Dynamic Panels: A Survey", Department of Economics, College Station, Texas A&Q University.

Banerjee, A., 1999, "Panel Data Unit Roots and Cointegration: An Overview", *Oxford Bulletin of Economics and Statistics 61*, 607-629.

Bilson, J.F.O., 1981, "The Speculative Efficiency Hypothesis", *Journal of Business 54(3)*, 435-51.

Civcir, I. 2002, "The Long-Run Volatility of Monetary Exchange Rate Model for a High Inflation Country and Misalignment: The Case of Turkey", *Russian and East European Finance and Trade*, forthcoming.

Coakley, J. and Fuertes, A. M., 1997, "New Panel Unit Root Tests of PPP", *Economic Letters 57(1)*, 17-22.

Cushman, D. O., 2000, "The Failure of the Monetary Exchange Rate Model for Canadian-US Dollar", *Canadian Journal of Economics, 33(3)*, 591-603.

Davidson, J.E.H., Hendry, D.F., Srba, F. and Yeo, S., 1978, "Econometric Modelling of the Aggregate Time-Series Relationship Between Comsumer's Expenditures and Income in the United Kingdom", *Economic Journal 88*, 661-692.

Dickey, D. A. and Fuller, W. A., 1981, "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", *Econometric 49*,1057-1072.

Dornbusch, R., 1976,"Expectations and Exchange Rate Dynamics", *Journal of Political Economy 84*, 1161-1176.

Engle, R.F. and Granger, C.W. J., 1987, "Co-integration and Error Correction: Representation, Estimation and Testing", *Econometrica55*, 251-276.

Flood, R.P. and Rose, A.K., 1999,"Understanding Exchange Rate Volatility without the Contrivance of Macroeconomics", *Economic Journal 109*, 660-672.

Frankel, J.A., 1979,"On the Mark: A Theory of Floating Exchange Rate Based on Real Interest Rate Differentials", *American Economic Review 69(4)*, 610-622.

Frankel, J. A. and Rose, A.K., 1996, “A Panel Project on Purchasing Power Parity: Mean Reversion within and Between Countries”, *Journal of International Economics 40*, 209-224.

Frenkel, J.A., 1976, “A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence”, *Scandinavian Journal of Economics 78*, 200-224.

Groen, J.J.J., 2000,“The Monetary Exchange Rate Model as a Long-run Phenomenon”, *Journal of International Economics 52*, 299-319.

Hall, A., 1990, “Testing for a Unit Root in Time Series with Pretest Data Based Model Selection”, *Working Paper*, North Caroline State University.

Husted, S. and MacDonald, R., 1998, “Monetary-Based Models of the Exchange Rate: A Panel Perspective”, *Journal of International Financial Markets, Institutions and Money 8*, 1-19.

Im, K.S. and Pesaran, M. and Smith, R., 1996, “Dynamic Linear Models for Heterogeneous Panels”, in Matyas, L. and Levestre, P. eds., The Econometrics of Panel Data: A Handbook of the Theory with Applications, Kluwer Academic Publishers, Boston, MA, 145-195.

Im, K.S. and Pesaran, M.H. and Shin, Y., 2003,“Testing for Unity Root in Heterogeneous Panels”, *Journal of Econometrics 115*, 53-74.

Johansen, S., 1991, “Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models, *Econometrica, 59(6)*, 1551-1580.

Johnson, H.G., 1977, “The Monetary Approach to the Balance of Payments”, *Journal of International Economics 7(3)*, 251-268.

Kao, C., 1999,“Spurious Regressions and Residual-Based Tests for Cointegration in Panel Data”, *Journal of Econometrics 90*, 1-44.

Levin, A. and Lin, C. F. and Chu, Chieng-Fu, 2002, “Unit Root Test in Panel Data: Asymptotic and Finite Sample Properties”, *Journal of Econometrics 108*, 1-24.

Levin, A. and Lin, C.F., 1993, “Unit Root Tests in Panel Data: New Results”, Department of Economics, *Discussion Paper 93-56*, University of California at San Diego.

Lothian, J.R., 1996,“Multi-Country Evidence on the Behavior of Purchasing Power Parity under the Current Float”, *Journal of International Money and Finance 16*, 19-35.

MacDonald, R., 1996, “Panel Unit Root Tests and Real Exchange Rates”, *Economics Letters 50(1)*, 7-11.

MacDonald, R., Cuaresma, J.C. and Fidrmuc, J., 2003, “The Monetary Approach to Exchange Rates: A Panel Evidence for Selected CEECs”, *Focus in Transition IMF 2*, 138-151.

Mark, N.C. and Sul, D., 2001, “Nominal Exchange Rates and Monetary Fundamentals Evidence form a Small Post-Bretton Woods panel, *Journal of International Economics 53*, 29-52.

Meese, R. A. and Rogoff, K., 1983, “Empirical Exchange Rate Models of the Seventies: Do they fit out of sample?”, *Journal of International Economics 14(2*), 3-24.

Moosa, I.A.,2000, “A Structural Time Series Test of the Monetary Model of Exchange Rates under German Hyperinflation, *Journal of International Financial Markets, Institutions and Money 10*, 213-223.

Mussa, M., 1976, “The Exchange Rate, The Balance of Payments, and The Monetary and Fiscal Policy under a Regime of Controlled Floating, *Scandinavian Journal of Economics 78*, 229-248.

O’Connell, P., 1998,“Market Frictions and Real Exchange Rates”, *Journal of International Money and Finance 15*, 405-418.

Oh, K. Y., 1996, “Purchasing Power Parity and Unit Root Tests Using Panel Data”, *Journal of International Money and Finance 15*, 405-418.

Papell, D. and Theodoridis, H., 1998,“Increasing Evidence of Purchasing Power Parity over the Current Float”, *Journal of International Money and Finance 17*, 41-50.

Papell, D., 1997, “Searching for Stationarity: Purchasing Power Parity under the Current Float”, *Journal of International Economics 43*, 313-332.

Pedroni, P., 1995, “Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the PPP Hypothesis”, *Indiana University Working Papers on Economics*.

Pedroni, P., 1997, “Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the PPP Hypothesis, New Results”, *Indiana University Working Papers on Economics*.

Pedroni, P., 1999, “Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors”, *Oxford Bulletin of Economics and Statistics, Special Issue*, 653-670.

Pesaran, M. H. and Shin, Y. and Smith, R., 1999, "Pooled Mean Group Estimation of Dynamic Heterogeneous Panels", *Journal of the American Statistical Association 94*, 621-634.

Pesaran, M.H., 2003,"A Simple Panel Unit Root Test in the Presence of Cross Section Dependence", *Cambridge Working Papers in Economics 346*, University of Cambridge.

Pesaran, M.H. and Smith, R., 1995,"Estimating Long-run Relationships from Dynamic Heterogeneous Panels", *Journal of Econometrics 68*, 79-113.

Phillips, P. and Perron, P., 1988,"Testing for a Unit Root in Time Series Regression", *Biometrica 75*, 335-346.

Phillips, P.C.B. and Moon, H.R., 1999,"Linear Regression Limit Theory for Nonstationary Panel Data", *Econometrica 67*, 1057-1111.

Psaradakis, Z. and Sola, M. and Spagnolo, F., 2004,"On Markov Error-Correction Models with an Application to Stock Prices and Dividends", *Journal of Applied Econometrics 19*, 69-88.

Rapach, D.E. and Wohar, M.E., 2004, "Testing the Monetary Model of Exchange Rate Determination: A Closer Look at Panels", *Journal of International Money and Finance, 23*, 867-895.

Reinhart, C.M. and Rogoff, K.S., 2002, "The Modern History of Exchange Rate Arrangements: A Reinterpretation", *NBER Working Paper 8963*.

Robertson, D. Symons, J., 1992,"Some Strange Properties of Panel Data Estimators", *Journal of Applied Econometrics 7*, 175-189.

Rogoff, K., 1996, "The Purchasing Power Parity Puzzle", *Journal of Economic Literature 34*, 647-668.

Rogoff, K., 1999, "Monetary Models of Dollar/Yen/Euro Nominal Exchange Rates: Dead or Undead?", *Economic Journal 109*, 655-659.

Ru-Lin Chiu, 2002, "Testing the Purchasing Power Parity in Panel Data", *International Review of Economics & Finance, 11(4)*, 349-362

Saikkonen, P., 1991, "Asymptocally Efficient Estimation of Cointegrating Regressions", *Econometric Theory 7(1)*, 1-21.

Sarno, L. and Taylor, M.P., 2002,"Exchange Rate Determination: Theories and Evidence", in The Economics of Exchange Rates, Cambridge University Press, Cambridge, UK, 1st edition.