UNIVERSITY OF KENT
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Essays on Exchange Rate Determination: An Analysis of Industrialised and Emerging Markets

by

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**TO MY PARENTS (IN MEMORIAM)
FOR LOVE AND HONESTY.**

DECLARATION

I hereby certify that the work embodied in this thesis is the result of my own investigations except where reference has been made to published literature.

I declare that this work has not already been accepted in substance, nor is it currently being submitted for any other degree.

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ABSTRACT

The main contribution of this thesis is to assess the importance of economic fundamentals in affecting the nominal exchange rate behaviour for a sample of industrialised and emerging market economies. Industrialised and emerging markets differ in terms of economic stability conditions that may affect distinctly the performance of traditional exchange rate models. The thesis consists of investigating the effects of three interrelated determinants on the exchange rate variability which are examined from the early 1980s until the early 2000s: the effect of monetary fundamentals, the effect of rational speculative bubbles and the effect of foreign debt.

Firstly, we test the monetary model for both types of market economies by making use of panel techniques that allow for a high degree of heterogeneity across countries. We find partial support for the monetary model for industrialised market economies but not for emerging ones. This constitutes a puzzle as we would expect countries with greater monetary instability to show a stronger association between exchange rates and monetary fundamentals. Secondly, we test for the presence of a rational speculative bubble driving the stochastic process of exchange rates away from the equilibrium level defined by monetary fundamentals. Our findings reveal that the hypothesis of periodically collapsing bubbles driving the exchange rate away from the fundamentals solution cannot be accepted for a sample of four industrialised market economy countries. Moreover, the results also revealed significant non-linearities and different regimes. The importance of these findings suggests that linear monetary models may not be appropriate to examine exchange rate movements. Finally, we investigate the effects of foreign debts on nominal exchange rate volatility for a sample of industrialised and emerging market economies where the level of foreign indebtedness and the access to international credit markets differ substantially. We also test for the use of monetary policy and international reserves to stabilise potential exchange rate volatility. Our findings confirm that foreign debts do generate effects on exchange rate volatility in both types of market economies, with more significant impacts on financially fragile economies (emerging markets). The results also revealed that monetary authority interventions in the foreign exchange market are a common practice in both categories of economies. The importance of these findings commends government authorities to keep foreign debt at levels consistent with macroeconomic stability.

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CHAPTER I

Introduction

The volatile behaviour of exchange rates provides ample evidence that in a world of liberalised financial flows, small transaction costs and large funds of mobile capital, and consequently huge pressure on financial markets, the so-desirable exchange rate stabilisation is an extremely difficult task for economists and policymakers. Despite such volatility and modelling difficulties there is still a strong belief in the monetary approach to exchange rate determination.

A brief historical summary leads initially to Friedman (1953), for instance, who argued that the instability in exchange rates is a consequence of instability in underlying economic structure (fundamentals). Primarily, this argument is based on the idea that if there is exchange rate instability it is due to disequilibrium in economic fundamentals. In an influential paper, Frenkel (1976) demonstrates strong support for the flexible-price monetary model for the German Mark/US dollar exchange rate during the German Hyperinflation of the 1920s. Bilson (1981) also estimates a model for major exchange rates during the 1970s float and finds broad support for the flexible-price monetary model. Driskell (1981), on the other hand, modelling the Swiss/US dollar exchange rate reports results broadly favourable to the sticky price monetary model in accordance with the Dornbusch (1976) story.

Despite these early results and supportive arguments in favour of a monetary theory of exchange rate determination, it could be argued that prior to the Bretton Woods' breakdown in the early 1970s, the consensus seemed to support the existence of a fairly stable exchange rate. The prevailing orthodoxy of the early 1970s, largely associated with the monetary approach to the exchange rate, assumed the much stronger proposition of continuous purchasing power parity (PPP). In the second half of the 1970s, this extreme position was abandoned due to the very high observed variability of exchange rates after the collapse of Bretton Woods. Studies published subsequently, mostly in the 1980s, could not refute the hypothesis of random walk behaviour in exchange rates, and reduced further the confidence in continuous purchasing power parity. This context led to the rather widespread belief that PPP was of little use empirically and that exchange rate movements were highly persistent and volatile (see Dornbusch, 1988).

A seminal paper by Meese and Rogoff (1983) compares the out-of-sample forecasts produced by various exchange rate models with forecasts produced by a random walk model and analyses their relative performance. They generated a succession of out-of-sample forecasts for each model and for various time horizons. The conclusion of this study was that none of the exchange rate models outperformed the simple random walk model. More recently, researchers have tested for cointegration between the nominal exchange rate and relative price levels, interpreted as testing for long-run purchasing power parity, which is equivalent to testing for mean reversion or stationarity in real exchange rates. Although some studies have found supportive evidence of reversion towards purchasing power parity for interwar period floats (see McNown and Wallace, 1989), other studies using cointegration methods have rejected this hypothesis even in high inflation economies. The rejection of the hypothesis of a mean reversion of the exchange rate towards purchasing power parity constitutes a significant criticism of the monetary approach.

Nowadays many researchers have concluded that there is significant exchange rates disconnect which is related to the remarkably weak short-term feedback links between the exchange rate and macroeconomic fundamentals. Rogoff (1996) has described it as a broader class of puzzles in international macroeconomics. He has also argued that to understand exchange rate behaviour one ultimately needs a broader model that explains the high stability that has been observed in asset markets. It is also true that the links between the exchange rate and the real economy are much more direct than for the stock market. In most economies, the exchange rate is the single most important relative price, one that potentially feeds back immediately into a large range of transactions [see Obstfeld and Rogoff (2000b)].

Obstfeld and Rogoff (2000b) also argue that exchange rates are remarkably volatile relatively to any model's underlying fundamentals such as interest rates, output, and money supplies; and no model seems to be good enough at explaining exchange rates even *ex*-post. According to them, although a range of goods is non-traded, there is always a broader range of goods that are traded and should tie down the exchange rate. In practice, markets for most traded goods are not fully integrated, and segmentation due to various transaction costs may be significant.

Many empirical studies have considered the interference of extraneous elements which affect the long-run equilibrium exchange rate. Evans (1991) refers to these elements as speculative, rational bubbles interfering with investors' expectations and causing substantial volatility in foreign exchange markets. Some early studies have supported the presence of rational bubbles in exchange rates. Meese (1986), for instance, provides strong evidence that the abrupt rise of the US dollar in 1984-1985, and subsequent drop between 1985 and 1988, was a direct consequence of bubbles driving exchange rate movements. Frankel and Froot (1990) reinforce this argument, investigating the high US dollar appreciation that occurred between 1984 and 1985, and show that standard observable macroeconomic variables (output, money growth rates, trade deficit) are not capable of explaining, and predicting *ex*-ante, the majority of short-term changes in exchange rates. According to them, the 1984-1985 episode was due to the existence of speculative rational bubbles in which market participants did not agree on the model for forecasting the exchange rate.

The impact of financial variables is another topic that has grown in importance in the literature to explain exchange rate behaviour. These variables are related closely to debt components that may cause a generalised fear in the market due to the possibility of defaults and so affect adversely exchange rates. More recently Devereux and Lane (2003) found evidence from a cross-section study that debt variables have a large influence on exchange rates. Financial fragilities associated with credit constraints in the international financial market may cause capital outflows and be an additional source of exchange rate instability. This instability comes from default risks by virtue of excess stocks of debts denominated in one currency only without an appropriate asset coverage. This research field is topical and will provide one theme for this thesis.

The complexity of exchange rate behaviour does not allow for robust conclusions as yet and motivates much current research. The development of new econometric approaches, alongside improvements in theoretical frameworks, may produce a better understanding of exchange rate determination. These advances also open up a number of new lines of enquiry, and create a challenging environment for researchers. For instance, the key hypothesis of the monetary approach to the exchange rate is that volatility in economic fundamentals is the main source of exchange rate variability.

In contrast to industrialised market economies, emerging market economies are historically more likely to demonstrate weak economic fundamentals such as an ever-expanding fiscal deficit, high inflation rates, low rates of economic growth and excess supply of money, etc., increasing risk and instability. The theoretical literature points out that such mismanagement of economic aggregates should produce direct effects on the exchange rate. A central contribution of this thesis is to examine this controversial debate in a comparative analysis of industrialised and emerging market economies. The approach is to select a sample of industrialised and emerging market economies and to investigate the determination of exchange rates using new econometric developments. The objective is to provide an additional contribution to the debate on exchange rate determination. To achieve this, the thesis will investigate if the hypotheses of macroeconomic fundamentals, of speculative rational bubbles, of foreign debts, are able to provide the alternative, compelling explanation for the volatile behaviour of exchange rates in industrialised and emerging market economies.

Four core chapters are developed and are briefly described as follows:

Chapter II reviews briefly previous and relevant empirical studies of exchange rate determination and provides a critical appraisal. It also provides an overview of each topic investigated in the thesis.

Chapter III investigates the effect of economic fundamentals on exchange rate behaviour for a sample of industrialised and emerging market economies. The approach consists of examining the existence of long-run relationships between nominal exchange rates and the traditional monetary approach using panel data techniques. Important contributions are offered in terms of a comparative study between different types of market economy. The analysis is executed through recently developed procedures for the detection of unit roots as well as for the identification of cointegrating properties in panel data. Finally, a pooled mean group estimation procedure is used which allows for wide heterogeneity across countries.

Chapter IV examines whether rational bubbles drive the exchange rate away from its fundamental equilibrium. A speculative rational bubble in the exchange rate is characterised by an explosive path. Explosive behaviour leads the exchange rate to diverge from the equilibrium level defined by monetary fundamentals. The approach of this chapter is to investigate the hypothesis of a periodically collapsing bubble governing exchange rate movements. Periodically collapsing bubbles are a special type of bubble, characterised as a stochastic process in which there are continuous periods of expansion and collapse. The chapter follows the methodology based on unit root and cointegration tests, and introduces a Markov-switching regime methodology in order to allow for a more robust analysis of this type of bubble.

Chapter V investigates the impact of foreign debts on exchange rate volatility using vector autoregression -VAR modelling. The investigation is conducted by applying this econometric method to a sample of industrialised and emerging market economies which have high and low levels of foreign indebtedness, respectively. The hypothesis tested is that changes in foreign debt impact on exchange rate volatility. It is expected that such an impact will differ across different market economies. The chapter also examines the possibility of government intervention, using either monetary policy or international reserves management, to stabilise exchange rate volatility.

CHAPTER II

LITERATURE SURVEY

Introduction

Standard economic theory states that exchange rate movements should be explained by a simple macroeconomic model made of money supply, output, and interest rates. Even though numerous studies have been produced over time, in practice fundamental variables have not proved to be completely helpful in explaining exchange rate variability. In particular, some results have revealed a significant exchange rate disconnect with its economic fundamentals which has become one of the most important puzzles in open economy macroeconomics. The efforts towards finding a solution to this controversy remain unresolved.

According to MacDonald (1999), the traditional literature has classified the investigations about the relationship between the exchange rate and its economic fundamentals into three different debates. The first focuses on the convergence of this relationship to the long-run equilibrium. The second is related closely to Meese and Rogoff (1983), and examines the power of fundamentals in generating accurate out-of-sample forecasts compared to a simple random walk model. The third analyses the impact of real and nominal shocks of fundamentals on the exchange rate volatility.

Some empirical investigations of exchange rates generate acceptable results when handling within-sample analyses, but they usually fail to impress when applied to out-of-sample forecasts. This is the strong conclusion revealed by the seminal work of Meese and Rogoff (1983). Nevertheless, it is a controversial conclusion as many finding over two decades have reported results in support of the monetary model. These studies are characterised by unit root and cointegration approaches using long spans of data. In fact, the classical monetary model relies heavily on the continuous PPP condition which requires long spans of data to hold. As Rogoff (1996) emphasised, estimates show that the speed of convergence to the PPP condition is low so that the exchange rate does not converge instantaneously to the long-run equilibrium defined by the model. However, this explanation seems not to be sufficient as even exchange rate models based on sticky-prices do not perform well.

The idea of long spans, nonetheless, may also suggest that short-run exchange rate movements would remain unexplainable given the standard models in use. It appears that further attempts to provide explanations about short-run exchange rate movements based solely on traditional macroeconomic fundamentals have not proved to be entirely successful. Indeed, one of the difficulties in improving on a simple random walk model, as argued by Meese and Rogoff (1983), may be related to tests based on short-run horizon data sets and heterogeneity across units which reduce the power of tests [see Rogoff, (1996), Pesaran *et al.* (1999)]. Empirical evidence shows that exchange rate deviations from economic fundamentals involve problems of data horizon, and country heterogeneity in cross-section studies, which may not reliably reject the hypothesis of unit roots. Thus standard monetary models would just be good enough to predict exchange rate behaviour in the long run and in a homogeneous context, rather than in the short run and in a heterogeneous one. These circumstances require that the empirical method assumes as important a role as the theoretical model itself. Chapter III uses an econometric approach that addresses partly this difficulty by decomposing the heterogeneity effects of economic fundamentals on exchange rate movements into short-run and long-run components across countries. In essence, the approach assumes heterogeneity across countries in the short run and homogeneity in the long run for a sample of industrialised and emerging countries.

Traditional macroeconomic fundamentals models are undeniably important in setting the parameters within which the exchange rate moves, but they do really not appear to tell the whole story. Taylor (1995) states, for instance, that even though some empirical studies have provided support for long-run relationships between the exchange rate and economic fundamentals, greater progress might be made if efforts were concentrated on the qualitative analysis of exchange rate determinants. This statement is related to problems of model misspecifications which can be crucial in investigations of the exchange rate and its fundamentals.

In fact, the debate about exchange rate models is still unclear and involves issues such as the most appropriate variables and price regime, sticky-price or flexible-price regimes. Thus, additional alternative explanations of exchange rate movements arise. For instance, the interference of extraneous elements may take the exchange rate away from its long-run equilibrium and may be very important. The literature refers to these interferences, such as the effect of news, the presence of bubbles[1], or even changes in portfolio preferences, as anomalies to the monetary model and suggests that they may drive the stochastic process of the exchange rate [see DeGenaro and Shrieves, (1997), Almeida *et al.* (1998), Hall *et al.* (1999), Psaradakis *et al.* (2001), Fornari *et al.* (2002), Chang and Taylor (2003), Oberlechner and Hocking (2004)]. Essentially, the studies based on these alternative scenarios investigate these anomalies as being inherently associated with

[1] There are different concepts of bubbles. Flood and Garber (1980) introduced a deterministic self-fulfilling concept of bubbles which expand through time at ever-accelerating rate. Blanchard and Watson (1982) developed the concept of stochastic rational bubbles which are stochastic processes completely independent of fundamentals. Froot and Obstfeld (1991) introduced the concept of intrinsic bubbles that are driven exclusively by nonlinearities in fundamentals. Evans (1991) presented the concept of periodically collapsing rational bubbles that expand and sudden collapse with probabilities π and $1-\pi$, respectively.

deviations of the exchange rate from its economic fundamentals. In practice, the anomalies act as a "contaminating virus" that violates the efficient market hypothesis and invalidates traditional monetary models of the exchange rate[2]. Significant efforts have been made to improve econometric techniques in order to detect accurately the presence of these extraneous events; however, there is still much to do. Chapter IV of this thesis investigates the presence of bubbles in exchange rates based on Evans (1991)'s concept, and on a Markov-switching regimes model. Evans (1991)'s bubble concept is appropriate to the reality of the foreign exchange market and to switching regimes models by modelling a bubble as a stochastic process that expands and collapses periodically.

Other "noises" to the fundamentals-exchange rate relationship, generically known as market frictions such as transaction costs, are also themes of the exchange rate literature (see Obstfeld and Rogoff, 2000). Moreover, efforts have also focused on investigating theoretical models which explicitly, or implicitly, propose a non-linear relationship between the exchange rate and fundamentals [see Meese and Rose (1991), Taylor and Peel (2000), Yue and Kana (2000), Taylor and Peel and Sarno (2001), Kilian and Taylor (2003)]. Krugman (1991) derived a target zone model for managed exchange rate regimes which establishes an S-shaped, non-linear fundamentals-exchange rate relationship. Krugman and Miller (1993) extended this study by deriving a non-linear theoretical model for free-float exchange rate regimes based on micro-foundations. Non-linear behaviour of exchange rates can also be attributed to non-linearities in government economic

[2] The efficient market hypothesis implies market agents with rational expectations and risk neutrality. It implies more than that: perfect information, sufficiently deep market (enough number of traders), perfect competition.

policies that may determine speculative attacks in the foreign exchange market (see Flood and Marion, 1998)[3].

However, the studies on exchange rate behaviour go much further and additional research efforts have focussed on two other fields. The first one is related to more purely financial models of exchange rate volatility which, conventionally, have been termed as market microstructure models in the economic literature. This emerging field consists of explanations of exchange rate movements based not on predictions about economic fundamentals, but essentially on the identification of supposedly recurring patterns in exchange rate behaviour. It is also usually called technical or chart analysis. In essence, this research proposes that the explanation of anomalous exchange rate movements comes from "information" exchanged by foreign exchange market dealers. Questionnaire surveys concluded that a significant proportion of market traders employ technical or chartist analysis [see Taylor and Allen (1992), Yin-Wong and Yuk-Pang (2000)]. In fact, studies based on microstructure models reveal that the dispersion of opinions among market participants works against the rational expectation and risk neutral hypotheses assumed by traditional exchange rate models [see O'Hara (1995), Fiers and MacDonald (2002), Dominguez (2003), Hashimoto (2005),].

The second group of emerging research is concerned with investigating exchange rate determinants in a context of dynamic general equilibrium. It introduces explicitly micro-foundations, nominal rigidities and imperfect competition into the analysis to derive the equilibrium exchange rate. These studies are usually termed New Open-Economy Macroeconomics. Although much of the

[3] In essence, Flood and Marion (1998) focus their study closely on fixed exchange rate regimes. For floating exchange rate regimes, most non-linear models use transactions costs or some other frictions, or still on a mixture of chartists and fundamentalists acting in the market.

literature associated with this group takes the redux model as a reference (see Obstfeld and Rogoff, 1995a), the precursor study was developed by Svensson and Wijnbergen (1989). The baseline of the redux model, and a significant part of the subsequent literature about this topic, investigates the international transmission mechanisms and the endogenous determination of interest rates and other asset prices, among them the exchange rate in an imperfectly competitive market (see Lane, 2001). Basically, they analyse the optimising behaviour of a representative agent and the impact of a monetary shock on real money balances, on output and on exchange rate behaviour. This research field goes beyond the scope of this thesis.

Finally, an important area of research has devoted attention to impacts of financial variables on exchange rate variability. These studies are classified into the third regime of research as proposed by MacDonald (1999), and basically investigate the importance of financial-economic equilibrium for exchange rate stability. Obstfeld and Rogoff (1995b) examine the impact of net foreign asset positions on exchange rate changes. De Gregorio *et al.* (1994) and Alesina and Perotti (1995) study the effects of fiscal policies on exchange rate determination[4]. Devereux and Lane (2003) use a cross-section model to analyse the impact of an optimal currency area and financial variables on exchange rate volatility. Chapter V of this thesis use a time-series approach to investigate the effect of high levels of net foreign debts on exchange rate volatility based on the ever-increasing default risk assumption. The basic idea is that high stocks of foreign debts lead to increases in default risks producing adverse effects on exchange rate volatility.

[4] Note that Obstfeld and Rogoff (1995b), De Gregorio et al. (1994) and Alesina and Perotti (1995) study equilibrium real exchange rates, not exactly exchange rate variability.

The following sections summarise the current debate about the relationship between the exchange rate and economic fundamentals. Section 2.1 consists of brief descriptions about some selected findings of the relationship between the exchange rate and standard economic fundamentals. Section 2.2 summarises some findings about bubbles in the exchange rate process. Section 2.3 outlines briefly the effect of foreign debts on the exchange rate. Section 3 summarises the main issues developed in the chapter.

2 – Literature Summary

2.1 – Effects of Economic Fundamentals on Exchanges Rates

The basic fundamentals-based exchange rate model is based on an efficient speculative market. Prices should fully reflect information available to market participants and it should be theoretically impossible for a trader to make excess returns from speculation. Participants are in general assumed to be risk neutral and have rational expectations. Under rational expectations and risk neutrality, the expected change in the exchange rate should differ from the actual change only by a rational expectation error which normally is white noise.

The underlying idea is that the exchange rate is a financial asset price and expectations about the future path of its determinants will be important for the current determination of the exchange rate. The traditional flexible price monetary model of exchange rates explains how future rational expectations about economic fundamentals can determine the variability in exchange rates. Starting from the purchasing power parity - PPP condition, the exchange rate, s_t, is determined by the following equation:

$$s_t = p_t - p_t^* \qquad (2.1.1)$$

where p_t and p_t^* denote the domestic price and the foreign price levels, respectively[5].

Provided that asset holders are able to adjust their portfolio instantaneously as bonds are assumed to be perfect substitutes, and financial capital flows are perfectly mobile, so the uncovered interest rate parity - UIRP condition holds. Hence, the expected change in exchange rate, denoted by $\Delta^e s_{t+1}$ ($s_{t+1}^e - s_t$), is determined by the interest rate differential:

$$\Delta^e s_{t+1} = \beta(i_t - i_t^*) \qquad (2.1.2)$$

where i is the domestic interest rate and i^* is the foreign interest rate. The expression (2.12) is the uncovered interest rate parity (UIRP) condition.

The nominal demand for money, denoted by m, is assumed to depend on real income, y, the price level, p, and the level of the nominal interest rate, i, so the equilibrium in the domestic and foreign money market is defined as:

$$m_t = p_t + \alpha y_t - \beta i_t \qquad (2.1.3)$$

$$m_t^* = p_t^* + \alpha^* y_t^* - \beta^* i_t^* \qquad (2.1.4)$$

Subtracting equations (2.1.3) and (2.1.4) so that the PPP condition in (2.1.1) holds, and assuming that there is no significant difference in income and interest elasticities of money demand

[5] The asterisk denotes a foreign variable.

among different countries ($\alpha = \alpha^*$ and $\beta = \beta^*$), then the exchange rate is determined as follows:

$$s_t = m_t - m_t^* - \alpha(y_t - y_t^*) + \beta(i_t - i_t^*) \qquad (2.1.5)$$

The monetary model expressed in (2.1.5) demonstrates that the exchange rate is a function of money supply differentials, real income differentials and interest rate differentials. Hence, an increase in domestic money supply, relative to the foreign money stock, will lead to a rise in exchange rate, s_t, i.e., a depreciation of the domestic currency in terms of foreign currency. Likewise, a rise in domestic real income will lead to an appreciation in the exchange rate.

Denoting the economic fundamentals as $z_t = m_t - m_t^* - \alpha(y_t - y_t^*)$ and assuming that the UIRP condition in (2.1.2) holds, the exchange rate can be rewritten as:

$$s_t = z_t + \beta\Delta^e s_{t+1} \qquad (2.1.6)$$

where $\Delta^e s_{t+1}$= ($s_{t+1}^e - s_t$) and,

$$s_t = (1+\beta)^{-1} z_t + \beta(1+\beta)^{-1} s_{t+1}^e \qquad (2.1.7)$$

Rational expectations is introduced by considering a full information set, denoted by Ω_t, so that s_{t+1}^e is defined by conditional expectations, $s_{t+1}^e = E(s_{t+1}/\Omega_t) = E_t s_{t+1}$. As long as market information is symmetric, the market participants are expected to know the process of exchange rate determination and its underlying stochastic structure:

$$s_t = (1+\beta)^{-1} z_t + \beta(1+\beta)^{-1} E_t s_{t+1} \qquad (2.1.8)$$

Using the law of iterated expectation and repeatedly substituting for $E_t s_{t+1}$ for n future time periods, the final form for the exchange rate is as follows:

$$s_t = (1+\beta)^{-1} \sum_{j=0}^{\infty} \left[\beta(1+\beta)^{-1}\right]^j E_t z_{t+j} \qquad (2.1.9)$$

The expression (2.1.9) states that the market participants expectations will depend on the underlying stochastic structure, defined by economic fundamentals, or government authority rules (z_t) defined for fundamentals at time t. The central idea is that the future policy (z_{t+j}) for fundamentals will determine current exchange rate behaviour. Any exchange rate deviation from its value determined by expected fundamentals in (2.1.9) may denote the presence of extraneous forces interfering with the convergence towards the equilibrium.

Rapach and Wohar (2004) applied panel tests to quarterly industrialised countries data using the monetary model of the exchange rate for the period 1973:1-1997:1. In essence, the study compared the performance of a country-by-country approach to a pooled data approach. The investigation analysed two aspects: (1) if the long-run parameter estimates met the values predicted by the theory, and (2) if there were cointegrating relationships between the exchange rate and the traditional economic fundamentals. Both aspects were examined for a country by country approach and a pooled data approach. The theoretical monetary model was as follows:

$$e_{it} = \beta_{i0} + \beta_{i1}\left(m_{it} - m_{it}^{*}\right) + \beta_{i2}\left(y_{it} - y_{it}^{*}\right) + u_{it} \qquad (2.1.10)$$

where $m_{it} - m_{it}^{*}$ and $y_{it} - y_{it}^{*}$ are the money supply differential and real income differential both between a home country and a foreign country (US is the reference country). The term u_{it} is assumed to be white noise, β_{i1}=1 and β_{i2}<0.

The long-run country-by-country parameters were estimated by ordinary least square-OLS, fully modified-OLS (FMOLS) due originally to Phillips and Hansen (1990) and in panel data to Pedroni (1996), dynamic-OLS (DOLS) due firstly to Saikkonen (1991) and Stock and Watson (1993) and in panel to Kao and Chiang (2000)[6]. The three estimators did not find the values hypothesised by the theory nor even the correct expected signs. Slightly better results were computed by using Johansen (1991) vector-error-correction maximum likelihood (JML) and Pesaran and Shin (1999) auto-regressive distributed lags (ARDL) so that some estimated signs met the theory. Next, country-by-country cointegration tests were applied by following Johansen (JML) and Engle-Granger (EG) approaches. The EG tests did reject the null hypothesis of no cointegration for the entire sample and JML tests only demonstrated a partial support for the alternative hypothesis of cointegration.

[6] FMOLS and DOLS estimators aim at eliminating the problems of asymptotic distribution biases as the cross-section dimension and heterogeneity rise, and endogeneity biases of feedback effects in multivariate regressions when the exogeneity requirement does not hold.

The panel investigation used the bias-corrected least squares dummy variable (LSDV) estimator developed by Kao and Chen (1995) and pooled mean group (PMG) estimator by Pesaran *et al.* (1999). These estimators differ from each other by considering parameter homogeneity (LSDV) and considerable parameter heterogeneity across units, respectively. The results for LSDV produced the correct signs for β_{i1} and β_{i2}, but β_{i1} was found to be statistically less than unity. The PMG method found the estimates statistically predicted by the theory. The panel cointegration tests were based on Pedroni (1995), Groen (2000) and Taylor and Sarno (1998) and found significant evidence of cointegration.

The final conclusion was that the better econometric performance associated with the pooled data approach in comparison to the country-by-country approach is due to two factors: (1) a significant increase in sample size when data are pooled; and (2) an increase in regressor variability for pooled data which improves the convergence to long-run parameters.

Kilian and Taylor (2003) investigated the non-linear behaviour in exchange rate by using a non-linear, exponential, smooth transition autoregressive (ESTAR) model for a sample of seven OECD countries for the period 1973:1-1998:4. The main idea is that economic models linking the exchange rate to its macroeconomic fundamentals do not take into account that the path to reach the long-run equilibrium may have a non-linear trajectory. Firstly they identified strong evidence of links between the quarterly exchange rate non-linearity and non-linear movements of relative prices[7]. Secondly, they also found that standard univariate tests of mean reversions have very low power to detect non-linear mean reversions which makes difficult to reject the null hypothesis

[7] This finding supports Rogoff (1996) concerning the very low speed of adjustment of real exchange rates.

of linear unit roots in real exchange rate behaviour. Thus, the use of an econometric approach that allows for ESTAR dynamics may improve significantly the long-run predictability.

The results revealed that the ESTAR model beat the random walk forecasts for six out of seven countries in the sample at long horizons. Hence, they concluded that the presence of ESTAR dynamics in the disequilibrium process facilitates the success of the random walk model in forecasting exchange rates. However, such success has more to do with non-linear dynamics than being a suitable explanation of exchange rate behaviour. Also, as formal statistical tests for random walk models have low power, the sample would have to include large fundamentals shocks to detect the presence of non-linear mean reversions[8]. This also explains why it is so difficult to beat a simple random walk model in empirical studies. However, at long horizons when larger departures from fundamentals take place, particularly in hyperinflation periods for instance, the exchange rate seems to behave according to economic theory[9].

Mark and Sul (2001) examine the long-run relationship between the nominal exchange rate and monetary fundamentals in quarterly panel data from 1973:1 to 1997:1 for nineteen countries. Similarly to Rapach and Wohar (2004), Mark and Sul also used the traditional monetary approach to the exchange rate in a panel data context. The study aimed, firstly, at investigating slope coefficient estimates in regressions of future exchange rate changes on monetary fundamentals. Secondly, it also aimed at examining the prediction accuracy of fundamentals-based regressions over the random walk model in out-of-sample forecasting. One of the

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[8] See also Rogoff (1999).

[9] Note that Bleaney and Mizen (1996) also concluded that random walk models are rejected against cubic form models and that the rate of mean-reversion for real exchange rates increases with distance from equilibrium.

hypotheses was that there is a tendency for regression slope coefficients and R^2 to increase in magnitude as the prediction horizon is extended. Furthermore, out-of-sample forecast accuracy of monetary fundamentals also tends to improve with prediction horizon.

In order to circumvent these two difficulties, Mark and Sul explored cross-sectional and time-series properties in a panel data set and imposed modest homogeneity restrictions across countries. By using panel-based tests of cointegration they focused, first, on whether nominal exchange rates were cointegrated with monetary fundamentals of a long-horizon regression. In fact, the test procedure consisted of testing for whether the slope coefficient is zero. It is equivalent to a test of the null hypothesis of no cointegration between the exchange rate and the fundamentals, but extended to the case of a panel data model.

The econometric model focused on panel estimation of the short-horizon predictive regression whose representation is as follows:

$$\Delta s_{it+1} = \beta x_{it} + e_{it+1} \qquad (2.1.11)$$

and,

$$e_{it+1} = \gamma_i + \theta_{t+1} + u_{it} \qquad (2.1.12)$$

where "i" indexes the country; x_{it} and s_{it} denote the deviation of the exchange rate from its fundamental value and the exchange rate of the country "i", respectively. The regression error e_{it+1} is an unobserved component, γ_i is an individual-specific effect, θ_t is a time-specific effect and u_{it} is supposed to be white noise.

The results found by Mark and Sul, using as a numeraire three different currencies, demonstrated that the asymptotic tests easily rejected the null hypothesis of no cointegration between the exchange rate and the monetary fundamentals at the 5% level, providing a robust result for the three numeraire currencies investigated. Relatively to the case of out-of-sample forecasting the analysis demonstrated that the coefficient on the monetary fundamentals was statistically significant and dominated the out-of-sample forecasting of a simple random walk model. This result refutes Meese and Rogoff (1983).

Groen (2000) estimated a panel version of a simple monetary model of the exchange rate. The data set consisted of quarterly data covering the period 1973:1-1994-4 for fourteen bilateral exchange rates taking as a benchmark both the US dollar and the Deutschmark. Basically, this study differs from Mark and Sul (2001) by using two different currency references in order to eliminate influences of US dollar instabilities. The Deutschmark was supposed to be more stable. Furthermore, Groen divided the sample into three different groups. The first group was made up of European Monetary System-EMS members, the second one of G7 members and the third one of G10 members.

The empirical long-run model for bilateral exchange rates is the classical monetary model as follows:

$$e_t = c + \left(m_t - m_t^*\right) - (1-\delta)\left(y_t - y_t^*\right) + \omega\left[E_t\left(e_{t+1}\right) - e_t\right] + \zeta_t \qquad (2.1.13)$$

$$\zeta_t = \omega_t + v_t - v_t^*$$

and rearranging (2.1.13) yields,

$$e_t = \beta_0 + \beta_1\left(m_t - m_t^*\right) - \beta_2\left(y_t - y_t^*\right) + u_t \qquad (2.1.14)$$

where m_t, m_t^* and y_t, y_t^* correspond to the logarithms of the domestic and international money supply, domestic and international real income, respectively. e_t is the logarithm of the nominal exchange rate, and v_t, v_t^* are white noise processes. The coefficients in (2.1.14) were expected to be $\beta_1 = 1$ *and* $\beta_2 < 0$.

The panel unit root test on the residuals of the panel regressions followed the method suggested by Levin and Lin (2002). The test for panel cointegration consisted of a two-step procedure on the panel data set. For the first step, the residuals were estimated by using the equation (2.1.14) with seasonal dummy variables included. The second step consisted of applying a specific version of the Augmented Dickey-Fuller test for panel data.

A significant contemporaneous error correlation in the exchange rate across countries was observed in the error matrix Ω_ε. Thus, the feasible generalised least squares estimator (FGLS) was employed in order to remove the cross-sectional dependence and also the cross-sectional heteroskesticity in the ε_{it}'s. This procedure aimed at recovering asymptotically normal properties for the *t*-ratios. Note that in contrast to Mark and Sul (2001) who used the panel dynamic OLS estimator, Groen (2000) used FGLS.

The tests based on a panel data framework and applied solely to EMS countries supported cointegration only when the Deutschmark was used as a benchmark. For G7 bilateral exchange rate countries consisting of the USA, Canada, France, Germany, Italy, Japan and the UK, the null hypothesis of no-cointegration could not be rejected. Nonetheless, when considering the other two groups, the G10 group consisting of G7 plus Switzerland, Sweden, Netherlands, and G10 plus four countries, respectively, the presence of cointegration was detected for both US and DM benchmarks.

In another related study, Lane (1999) tested the nominal exchange rate against the US dollar on the long-run inflation differential and the long-run changes in the real exchange rate. The study analysed the behaviour of the exchange rate in countries that experienced average annual depreciation rates in excess of 30 per cent, countries that demonstrated appreciations against the US dollar and thirty-one countries that did not have autonomy over monetary or exchange rate policy for the period 1974-1992. In essence, this investigation examined the long-run determinants of the nominal exchange rate based on cross section data.

Differently from the aforementioned studies, Lane used an alternative model of PPP (Purchasing Power Parity) to model the long-run movements of the exchange rate as depending on the long-run inflation differential and the long-run change in the real exchange rate. The approach followed the recent literature which has modelled long-run inflation using a Barro-Gordon style model[10]. The use of the long-run real exchange rate as a predictor of long-run movements in nominal exchange rates allows the analysis of additional effects from productivity gains and changes in the terms of trade. Particularly, it is closely related to Balassa-Samuelson effects[11].

[10] Barro-Gordon modelling for long-run inflation is linked to variables such as trade openness, country size, central bank independence, political stability and others.

[11] In essence, Balassa-Samuelson effect assumes productivity gains as determinant for real exchange rate movements [see Faria and León-Ledesma (2003)].

The basic empirical model was decomposed into two parts: the long-run inflation differential and the long-run change in the real exchange rate as follows:

$$\pi_i^e = \pi_i - \pi^* + \pi_i^R \qquad (2.1.15)$$

where π_i^e is the rate of nominal exchange rate depreciation, π_i is the rate of inflation in country "i", π^* is the inflation rate of the benchmark country and π_i^R is the rate of real exchange rate depreciation.

The expression (2.1.16) summarises the hypotheses tested by Lane:

$$\pi_i^e = \pi^e(z_i, g_i, \pi^{TTi}, \pi^*, g^*) \qquad (2.1.16)$$

where z_i may denote trade openness, size of country "i" and the degree of central bank independence, g_i and g^* are the home and foreign output growth rates, respectively, and π^{TTi} is the growth rate of the terms of trade.

The empirical results pointed out that both inflation and the real exchange rate were significant determinants of the rate of growth of the nominal exchange rate. Openness of the economy was found to be significant in explaining the rate of nominal depreciation. The evidence for central bank independence, country size, political stability, and past inflation was weaker. Although the output growth rate was important in determining the long-run nominal exchange rate movements, the terms of trade did not have a similar importance.

Flood and Rose (1995) also investigated the ability of a monetary model in explaining exchange rate volatility. The empirical model focused on bilateral US dollar exchange rates from 1960 through 1991 for OECD countries. Nevertheless, the central objective was to make a comparison between exchange rate volatility and their fundamentals in regimes of fixed and floating rates. It was expected that macroeconomic fundamentals presented higher volatility in fixed exchange rate regimes than in floating regimes. The study was based on simple regression equations of flexible price and sticky price regimes. This is an important difference compared to the earlier investigations in that the price regime is also assumed to be relevant in explaining the trade off between macroeconomic fundamentals and exchange rate volatility.

The starting point was to contest empirically the traditional belief that exchange rate stabilisations could only be reached at the expense of more volatility in other macroeconomic variables. The volatility of variables such as the money supply and output, for example, did not appear to be significantly different during regimes of fixed and floating exchange rates. Countries that chose not to manage their exchange rates by allowing a persistent volatility could intervene in their foreign exchange markets without having the volatility transferred into another macroeconomic variable.

The strategy consisted of using two different models: the first model was a traditional monetary framework which consisted of a structural model constructed in a flexible-price context; the second one is a more complex model which was defined in a sticky-price context. The flexible-price model and the sticky-price model are defined as in (2.1.17) and in (2.1.18), respectively:

$$e_t = \left(m - m^*\right)_t - \beta\left(y - y^*\right)_t + \alpha\left(i - i^*\right)_t - \left(\varepsilon - \varepsilon^*\right)_t - v_t \qquad (2.1.17)$$

and,

$$e_t = \left(m - m^*\right)_t - \beta\left(y - y^*\right)_t + \alpha\left(i - i^*\right)_t - \left(\varepsilon - \varepsilon^*\right)_t - \phi/\theta r_t - \phi/\theta^2 E_t\left(r_{t+1} - r_t\right)$$
$$- \theta^{-1} E_t\left[\left(e_{t+1} - e_t\right) + \left(p^*_{t+1} - p^*_t\right)\right] \qquad (2.1.18)$$

In the flexible-price model e_t, $\left(i - i^*\right)_t$, $\left(m - m^*\right)_t$, $\left(y - y^*\right)_t$ and $\left(\varepsilon - \varepsilon^*\right)_t$ denote the exchange rate, home and foreign interest rate differential, home and foreign money supply differential, home and foreign income level differential and the last term is the monetary shock. In the sticky price model, the variable $r_t = i_t - E_t\left(p_{t+1} - p_t\right)$ is the *ex*-ante real interest rate. The term $-\theta^{-1} E_t\left[\left(e_{t+1} - e_t\right) + \left(p^*_{t+1} - p^*_t\right)\right]$ in (2.1.18) constitutes a central point of the sticky-price framework as it denotes rejection of the classical assumption of continuous purchasing power parity conditions implicit in traditional monetary models with flexible prices.

The conclusion was that few macroeconomic variables for OECD countries experienced significant changes in volatility which coincided with differences in exchange rate regimes. Although the exchange rate volatility showed a considerable change in different regimes, it was not observed that such behaviour was followed by other macroeconomic variables. Hence, two important conclusions were drawn from this investigation: (1) non observance of any strong trade off between exchange rate volatility and the volatility of a variety of different macroeconomic variables; (2) as a result of (1), little evidence on this trade off was found, and thus to reduce exchange rate volatility does not imply transferring instability into other macroeconomic variables.

In summary, the results presented in this sample of recent papers, based on the classical monetary model of the exchange rate, address important questions of differences in estimators, reference currency, pricing regimes and non-linear behaviour. Nevertheless, the different conclusions keep the debate very much alive, and open to new contributions. For instance, one can observe that a common theme is the analysis of industrialised countries that, in general, are characterised by stable economic fundamentals and weak heterogeneity. Monetary models are supposed to offer better performance in environments of economic instability. Chapter III of this thesis will consider a standard monetary model of the exchange rate, but it essentially aims at capturing heterogeneous and homogeneous effects across countries using a panel data analysis of industrialised and emerging economy countries with distinct economic features.

2.2 - Effects of Rational Speculative Bubbles on Exchange Rates

The concept of a speculative bubble is often used to explain the reason why asset prices appear to take long shifts away from their "fundamental" values. The existence of speculative bubbles in financial markets has been a frequent topic of debate. In particular, bubbles in the foreign exchange market are related remarkably to substantial exchange rate volatility not followed by similar behaviour in fundamentals. This economic phenomenon has been an important characteristic of floating exchange rate regimes observed in the past thirty years.

The literature postulates that, on the basis of the hypothesis of efficient markets, an asset should be priced according to its market fundamentals. However, deviations of an asset price from the value dictated by its market fundamentals may also be interpreted as evidence of rationality. Actually, the existence of a stochastic process and the adoption of rational expectations help to clarify considerably the nature of a rational bubble. The concept of a rational bubble is associated with a well-defined stochastic process. Hence, rationality inherent in this stochastic process generates self-fulfilling behaviour among market participants that influences asset prices. In fact, the self-fulfilling behaviour is usually initiated from extraneous events to the market which may assume a speculative form and leads to a rational bubble trajectory. The increasing frequency of this scenario has led the academic community to focus significant attention on the problem.

Studies on exchange rate bubbles have produced mixed results. For example, West (1987) applies a volatility test to foreign exchange markets and finds no bubbles. He attributes only partially exchange rate variability to monetary shocks and deviations from the purchasing power parity condition. Meese (1986), however, reports evidence of bubbles for Dollar/Deutschmark and Dollar/Sterling exchange rates. Although the investigation of bubbles involves broader sets of models, the flexible-price monetary approach once again offers a convenient framework to derive a speculative bubble due to the emphasis placed on the role of the expected exchange rate.

The equilibrium exchange rate arising from (2.1.8) is the solution of a first order difference equation in expectaitons that admits one particular solution (2.1.9) and also an infinite number of solutions as follows:

$$s_t = s_t^n + b_t \qquad (2.2.1)$$

where s_t^n and b_t are the non-bubble solution (given only by expectations of economic fundamentals) and the speculative bubble solution, respectively. Essentially, only one solution corresponds to the fundamentals solution and the rest may be considered bubble solutions. The speculative bubble term (b_t) determines the non-fundamentals solution to the exchange rate if investor beliefs rely on it. For (2.2.1) to be a solution of (2.1.8) b_t must follow:

$$b_t = \beta\, E_t\, b_{t+i} \qquad (2.2.2)$$

Taking expectations of (2.2.1) and assuming (2.2.2) results in:

$$E_t\, s_{t+i} = E_t\, s_{t+i}^n + \beta^{-1} b_t \qquad (2.2.3)$$

Since β is assumed to be less than one, from (2.2.3) the expected value of b_t ($E_t\, b_{t+i}$) will explode as i approaches infinity (to $+\infty$ for b>0 and to $-\infty$ for b<0). Considering that the exchange rate is a relative price, and there is not any economic sense in having a negative price, b is always assumed to be greater than zero (see Hallwood and MacDonald, 1994). This implies that the exchange rate will vary by an amount equal to βb even when economic fundamentals "z" in (2.1.9) remain unchanged[12].

Nonetheless, the presence of a bubble driving a stochastic exchange rate process only makes sense if a suitable and precise market fundamentals model is used. An inappropriate definition of market fundamentals may lead to a misleading conclusion about the bubble hypothesis. One might incur large mistakes by concluding in favour of a bubble process when there is actually a specification error influencing the results. Without a clear definition of a precise market fundamentals model, it is impossible to isolate the time path which characterises the presence of a bubble.

The first formal academic test for detecting bubbles was conducted by Flood and Garber (1980) when they investigated the presence of bubbles in the price level during the German hyperinflation of the early 1920s. They found no significant evidence of bubbles and concluded that the data were consistent with Cagan's (1956) monetary model. However, their methodology had shortcomings by testing for only deterministic bubbles, that is, those that just grow at the rate of interest and explode as time passes. It was an unrealistic mode of analysis in this context as it is unreasonable that a financial asset price can just grow without limits in a real economy. Further, the tests presented a second shortcoming in that the regressor for the bubble parameter was an exponential function of time, leading it to explode as the number of observations became large and to degenerating effects on the asymptotic distribution of the test statistic.

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[12] Economically, this is an "open window" of eventual pure capital gains, unrelated to fundamentals, through which investors may make profits by arbitrating their funds in different currencies in the international foreign exchange market (see Sarno and Taylor, 2002).

West (1987) implemented Hausman's (1978) specification test to investigate the presence of bubbles in the US stock market for the period 1871-1980 and found significant evidence of bubbles. Moreover, Meese (1986) using the same test specification found strong evidence of bubbles for the Dollar/Deutschmark and Dollar/Sterling exchange rates. Nonetheless, the methodology applied by both Meese and West consisted of an approach that tested for a joint null hypothesis of no bubbles and model misspecification against the alternative in which the regression errors were correlated with the regressors. The problem arose from the fact that the alternative might also include misspecification problems such as structural model misspecification, market fundamentals misspecification and even the presence of bubbles or other types of irrational behaviour. The presence of any of these factors might lead a researcher to reject the joint null hypothesis, and mistakenly be interpreted as a bubble.

Another difficulty with many earlier studies about bubbles relates to the use of a deterministic bubble concept associated with standard unit root and cointegration tests. These tests have failed to detect a bubble even in cases when it is substantial both in magnitude and volatility. The explanation is that such tests are more likely to detect a bubble only in its expansion period. In fact, the literature has shown that traditional integration and co-integration tests are known to have low power for detecting a bubble collapsing periodically. Evans (1991) also demonstrated that standard unit root tests are only capable of detecting bubbles that continue to infinity. Actually, he introduced the concept of periodically collapsing rational bubbles which are stochastic and show periods of expansion followed by periods of collapses. This stochastic concept constituted an important advance in investigations of bubbles as it allowed for more robust results that were closer to reality.

Psaradakis *et al.* (2001) investigated the possibility of a rational bubble being present in the German hyperinflationary process. The study used monthly data for the wholesale price index and money supply for the period 1914:12 – 1924:12. They applied a simple test procedure to detect the presence of explosive bubbles that collapse periodically based on the class of non-stationary, varying coefficient autoregressive models with a stochastic unit root. The underlying idea consisted of, firstly, testing whether the time series under study are described as having a random root with unit mean and then to identify periods during which explosive behaviour is related to the presence of bubbles.

Their research assessed whether the value of the price index is more explosive than the relevant fundamentals variables. If so, the conclusion would be to recognise the presence of rational bubbles driving the temporal process. Hence, models that allow for random coefficient variation are likely to provide suitable representations of the dynamics of series influenced by bubble components. The test was based on the following random-coefficient autoregressive model:

$$X_t - \lambda_t - \sum_{t=1}^{p} \phi_1 X_{t-1} = \rho_t (X_{t-1} - \lambda_{t-1} - \sum_{t=1}^{p} \phi_i X_{t-i-1}) + \varepsilon_t, \qquad t = 1,\dots,T \qquad (2.2.5)$$

where $\lambda_t = \delta_0 + \delta_1 t + \delta_2 [t(t+1)/2]$. It is assumed that $\{\rho_t\}$ *i.i.d.* $(1, \omega^2)$; $\{\varepsilon_t\}$ *i.i.d.* $(0, \sigma_\varepsilon^2)$ independent of ρ_t and all the zeros of the polynomial $z^p - \phi_1 z^{p-1} - \dots - \phi_p$ appear inside the unit circle. The basic hypotheses are $H_o: \omega^2 = 0$ against $H_1: \omega^2 > 0$ which are equivalent to testing the stationary and non-stationary processes.

In essence, the study involved tests for the simultaneous presence of random unit roots in the price index and its underlying fundamentals. The conclusion was that the explosive behaviour in the price index was not accompanied by synchronous explosive behaviour in fundamentals. Thus, the hypothesis of the presence of rational bubbles driving the temporal process during the German hyperinflation could not be rejected. The standard behaviour of explosive bubbles is that the temporal series cannot be made stationary even in first difference. In essence, Psaradakis *et al.* (2001) followed a research line based on stationarity analysis with an empirical model comprising an autoregressive root that varies randomly around a unit mean, which occasionally exhibits explosive behaviour.

Hall *et al.* (1999) analysed the integration properties of the money supply, consumer price index and the exchange rate against the US dollar in Argentina. The data set consisted of 82 monthly observations from 1983:1 to 1989:11. This period was characterised by episodes of hyperinflation in Argentina. The objective was to investigate whether the non-stationarity of the consumer price index might be attributable to explosive rational bubbles. Although this research was also based on stationarity tests, Hall *et al.* contributed a new approach consisting of a generalisation of the ADF test procedure in conjunction with Markov-switching regressions (see Hamilton, 1994).

In fact, Hall *et al.* argued that the great problem for detecting a bubble governing a stochastic process lies in identifying and separating expanding periods from collapsing ones. The reasoning is that the existence of a rational bubble is consistent with temporal processes characterised by the presence of an explosive autoregressive root. An efficient test should allow for the possibility of changes in the dynamic behaviour of asset prices throughout the sample.

Thus, they proposed to change the standard ADF test procedure to allow for the possibility that different dynamic behaviour of a variable may take place in different parts of the sample. This can be framed in the specific context of periodically collapsing rational bubbles, present in separated regimes, demonstrating expanding and collapsing phases. In fact, this new specification for the ADF test has similarities with the test procedure proposed by Perron (1990) and Perron and Volgesang (1992) in which it is assumed that there is a one-time exogenous break in mean. The test proposed by Hall *et al.* is expressed in (2.2.6) with an observable indicator $s_t \in \{0,1\}$ so that:

$$\Delta y_t = \mu_0(1-s_t) + \mu_1 s_t + \left[\phi_o(1-s_t) + \phi_1 s_t\right] y_{t-1} + \left[\sum_{j=1}^{k} \psi_{0j}(1-s_t) + \psi_{1j} s_t\right] \Delta y_{t-1} + \sigma_e e_t \tag{2.2.6}$$

where $\{e_t\}$ is equal to $i.i.d.\{0,1\}$. The random sequence $\{s_t\}$ is specified to be a homogeneous Markov chain on the state space $\{0,1\}$ with probability:

$$\Pr\{s_t = 1 / s_{t-1} = 1\} = p$$
$$\Pr\{s_t = 0 / s_{t-1} = 1\} = 1 - p$$
$$\Pr\{s_t = 0 / s_{t-1} = 0\} = q$$
$$\Pr\{s_t = 1 / s_{t-1} = 0\} = 1 - q$$

The probability of the state (regime) at time t depends on the state at time $t-1$. The test of hypotheses within the Markov-switching ADF framework is based on the $t-ratio$ associated with the maximum likelihood estimates of ϕ_0 and ϕ_1. The condition $\phi_0 = \phi_1 = 0$ denotes the unit-root null hypothesis which if valid is inconsistent with the presence of rational bubbles. The one-sided alternative hypothesis, consistent with the existence of an explosive rational bubble in price should produce $\phi_0 < 0$ and $\phi_1 > 0$.

The results suggested that during the 1989 hyperinflation it was possible to observe a clear explosive regime in the consumer price index associated with the rapid growth in the money supply. The implication was that the 1989 hyperinflation in Argentina was related to negative fundamentals behaviour, and not to rational bubbles. Nonetheless, the explosive price in the period 1988:6 to 1988:8 was not associated with a similar explosive behaviour in neither the money supply nor the exchange rate which is a finding consistent with a rational bubble in the consumer price index.

Taylor and Peel (1998) investigated the possibility of having a rational bubble process in the US real stock price for the period 1871-1987. In contrast to the aforementioned studies in which the approach was based on unit root tests, they examined the presence of periodically collapsing bubbles using cointegration tests applied to a standard present value model of stock prices given by:

$$p_t - r^{-1} d_t = (1+r) r^{-1} \sum_{j=1}^{\infty} (1+r)^{-j} E_t \Delta d_{t+j} \qquad (2.2.7)$$

where p_t is the real stock price at time t, d_{t+j} is the real dividends paid to the shareholder. The term $(1+r)^{-j}$ represents the discount factor and E_t is the mathematical conditional expectation for information at time t. The unique solution for equation (2.2.7) depends on the transversality condition $\lim_{n\to\infty}(1+r)^{-n} E_t p_t = 0$, so that $p_t = f_t$. This implies that p_t and d_t are I(1), but cointegrated with cointegrating parameter r^{-1}. However, this condition may fail, and so a bubble process may take place. If there are bubbles, the right-hand side of equation (2.2.7) is augmented by B_t. As B_t is a nonstationary process, then p_t and d_t cannot be cointegrated in the presence of bubbles.

The cointegration test consisted of estimating, firstly, RALS Dickey-Fuller and standard Dickey-Fuller statistics, CR_{τ_a} and CR_τ, respectively at the 5% level[13]. The RALS statistic was obtained from a two-step OLS estimation procedure. The first step estimates a standard equation $y_t = \psi z_t + u_t$. The residuals ($\hat{u}_t$) and the variance ($\hat{\sigma}^2$) are used to derive the vector of covariates $\hat{w} = \left[\left(\hat{u}_t^3 - 3\hat{\sigma}^2\hat{u}_t\right)\left(\hat{u}_t^2 - \hat{\sigma}^2\right)\right]'$. The second step consists of estimating the variance of β^* by RALS from the following equation:

$$y_t = \mu + \beta^* z_t + \gamma \hat{w}_t + \eta_t \qquad (2.2.8)$$

with the covariate vector introduced. The RALS Dickey-Fuller statistic was defined as $\tau_a = \beta^* / V(\beta^*)^{1/2}$ where $V(\beta^*)$ is the variance of β^*[14]. The underlying idea was that a cointegration test based on an estimator originally designed to be robust to the presence of skewness and kurtosis may be biased towards incorrect rejection of

[13] Residual Augmented Least Square – RALS.
[14] Essentially, it is the t-ratio of β in (2.2.8).

non-cointegration when periodically collapsing bubbles are present. This biased rejection is more likely to occur when standard non-cointegration tests based on ordinary least squares are applied. Thus, the vector $\hat{w}=\left[\left(\hat{u}_t^3 - 3\hat{\sigma}^2\hat{u}_t\right)\left(\hat{u}_t^2 - \hat{\sigma}^2\right)\right]'$ aims at correcting the effects of skewness and kurtosis usually found in standard Dickey-Fuller tests under the influence of periodically collapsing rational bubbles driving the process.

Finally, the 5% critical values for RALS Dickey-Fuller ($CR_{\tau a}$) and standard Dickey-Fuller statistics (CR_{τ}) were generated. Next, the residuals from the p_t on d_t regression were tested using the $CR_{\tau a}$ and CR_{τ} statistics. The computations based on either $CR_{\tau a}$ or CR_{τ} rejected the null hypothesis of non-cointegration at the 5% level. This result means that the stock price series did not contain a bubble component.

Van Norden (1996) developed a new test for speculative bubbles in exchange rates and applied it to the bilateral exchange rate of the Japanese yen, German mark and Canadian dollar against the US dollar. The empirical work used different specifications of fundamental models for exchange rates and applied the test for the period 1977 to 1991. In particular, it is a new test in the sense that it gives more information about a particular form of bubble and differs from the others that have a more general approach to detection and analysis. The focus of the test is on a specific stochastic type of bubble which is supposed to grow and to collapse fully or partially. Assuming a certain probability and size of these collapses, it can be demonstrated that the specific behaviour of bubbles may lead to regime-switching behaviour in exchange rate innovations.

Van Norden used a switching regime model, arguing that it may offer new explanations for large, abrupt exchange rate movements which may be linked to movements in other macroeconomics variables. The theoretical model was as follows:

$$S_t^* = \sum_{j=0}^{T} a^j E_t\left(f\left(X_{t+j}\right)\right) \qquad (2.2.9)$$

where S_t^* is the fundamentals exchange rate, X_t is the set of macroeconomics variables and $E_t(f(X_{t+j}))$ is the current expected value of fundamentals at $t+j$. The correlation coefficient is $0 \leq a \leq 1$.

Equation (2.2.9) is the fundamental solution for the exchange rate, namely, it is the correlation solution between the exchange rate and macroeconomic fundamentals. However, it is not the only possibility as any other set of exchange rate values such that $S_t \neq S_t^*$ may be defined as having bubbles. S_t is the actual exchange rate determined by market forces. The size of the bubble, in turn, may be defined as:

$$b_t = s_t - s_t^* \qquad (2.2.10)$$

and if the bubble is expected to grow over time then $a < 1$ and:

$$b_t = a\, E_t(b_{t+1}) \qquad (2.2.11)$$

Two simple exchange rate approaches were used to test for the presence of bubbles. The first one was the traditional flexible-price monetary model based on the purchasing power parity

(PPP) condition and on the uncovered interest rate parity (UIRP) condition. The second approach was based on the fact that the real exchange rate is a rate which equilibrates the external sector of an economy, therefore deviations from this rate (b_t) should be a function of external sector imbalances. Thus, deviations from current account balances denote actual deviations from fundamentals which might account for depreciations in exchange rates. A final measure used the sticky-price monetary model in which the overshooting phenomena itself was assumed as evidence of bubbles.

In short, Van Norden found that the results were sensitive to changes in the definition of fundamentals for exchange rates. The evidence was not found to be robust: some results supported the presence of bubbles when using UIRP and overshooting models for the Canada/US exchange rate or a PPP model for the Japan/US exchange rate.

Although the switching regime models attenuate the problem of model misspecification by focussing on differences in regimes, the sensitivity of fundamentals-based exchange rate models obtained by Van Norden reveals once again the relevance of this issue in investigating the presence of bubbles. Furthermore, the study also assumed that any deviation from fundamentals was a proof of bubbles driving the stochastic process. However, recent evidence of nonlinear exchange rate behaviour may reject this hypothesis [see, for instance; Taylor, Peel and Sarno (2001), Kilian and Taylor (2003)].

Charemza (1996) analysed the collapse of the Polish foreign exchange market in 1989. In that time analysts believed that the collapse was caused by the bursting of a stochastic speculative bubble caused by increasing inflationary expectations after the collapse of communism and financial liberalisation. The investigation was based on a present value model of the exchange rate. Given that a bubble is not directly observable, the study used a method of simulating repeatedly an unobservable non-linear component, and then introduced it into an econometric model as an observable variable.

The hypothesis that inflation expectations in Poland caused the development and the bursting of a bubble in the Polish foreign exchange market was based on the following model:

$$p_t = \overline{\varphi}_1 \pi_t + \overline{\varphi}_2 v_t + \overline{\varphi}_3 b_t + u_t \qquad (2.2.12)$$

where p_t denotes the free market price of the US dollar, π_t is monthly domestic inflation, v_t is an uncertainty term (defined as the quadratic spread of dollar prices among five main local markets in Poland), and b_t is the simulated component from the ith replication of $b_t = \theta b_{t-1} + e_t$ [see also Pesaran and Pesaran (1993); Laroque and Salanie (1993)] .

The test procedures relied, firstly, on estimating equation (2.2.12) by OLS and FM-OLS methods without the presence of the bubble component, and then applying ADF tests to the residuals with four lag augmentations. The results revealed that the US dollar price in Poland (p_t), inflation (π_t), and the spread variable v_t, were integrated of order I(1). Secondly, equation (2.2.12) was estimated by OLS and FM-OLS methods with the simulated bubbles process b_t included, and once again the ADF tests on u_t were applied. The results provided evidence in favour of speculative bubbles in the Polish foreign exchange market during that period.

Charemza used, in essence, the unit root approach to analyse the presence of bubbles in exchange rates. But, as a bubble is an unobservable process, the conventional ADF test may mistakenly either reject or accept its presence. The use of a simulated component to denote the presence of bubbles aimed at eliminating biases in estimated parameters, caused by any unobservable element in the stochastic process, did not take into account that the bias may be attributed to model misspecifications or non-linear behaviours.

In general, the studies of bubbles briefly described in this review followed approaches based on either ADF or cointegration tests. Although these approaches may bring contributions to the detection of certain types of bubbles, they may also demonstrate some degree of inefficiency when dealing with rational bubbles collapsing periodically, and when the stochastic process is shown to behave non-linearly. Chapter IV of the thesis will address these issues by adopting the Markov-switching regime approach to investigate the presence of bubbles collapsing periodically. The problem of model misspecifications is more complex and goes beyond the scope of this study.

2.3 - Effects of Foreign Debts on Exchange Rates

Some studies have highlighted the effects of financial variables, especially foreign currency-denominated debts, on exchange rate variability. The existence of un-hedged foreign currency denominated debts have significant effects on the financial sector, on corporate balance sheets and on the public sector with potential unwelcome side effects on the exchange rate. A high level of foreign currency denominated debts introduces additional costs to exchange rate variability when it is not hedged by at least an equivalent amount and quality of financial assets. These side effects are expected to be more significant in emerging market economies than industrialised market economies by virtue of the credit constraints imposed by the international financial market.

Eichengreen and Hausmann (1999), for instance, state that international credit constraints, especially for developing countries, are related to distorting guarantees provided by governments and international multilateral entities. Excess credit availability allows borrower countries to assume excessive debts protected by guarantees. Nevertheless, financial fragilities emerge and credit constraints are imposed as a result of increasing default risks. Furthermore, there are difficulties of borrowing abroad long term in domestic currency. This raises the problem of currency mismatches or maturity mismatches as long term domestic investment projects are normally foreign currency-financed projects. Ultimately, Eichengreen and Hausmann (1999) conclude that many emerging market economies that own high stocks of debts may have huge difficulties to manage their exchange rate volatility. These difficulties affect the financial credibility of a country and may lead to a more volatile exchange rate.

Devereux and Lane (2003) also examined the effect of debts on the exchange rate and reinforce the argument that high foreign debt stocks have different impacts on exchange rate movements for developed and developing countries. Developed countries have free access to the international financial market so that they can borrow by issuing assets denominated in their own currencies. Developing countries, in turn, are conditioned to severe borrowing constraints in international financial markets and only issue debts in foreign currencies. Following an adverse external shock, the credit constraints may produce perverse effects on exchange rate volatility in highly indebted economies due to increases in default risk. Furthermore, the increase in default risk caused by a high stock of foreign debts has effects on exchange rate volatility by also affecting the risk premium associated with investments in foreign currency. The risk premium becomes more volatile by virtue of the financial fragility degree demonstrated by a highly indebted country.

The following argument presents the underlying idea about the impact of increased default risks on the risk premium and the collateral effects on the exchange rate.

$$\Delta_i^e s_{t+i} = \alpha + \beta\left(i_t - i_t^*\right) + v_{t+i} \qquad (2.3.1)$$

(2.3.1) is the uncovered interest rate parity condition in which $\Delta_i^e s_{t+i}$ denotes the expected change in exchange rate, $\left(i_t - i_t^*\right)$ is the interest rate differential between the home country and the rest of the world, and α is the risk premium.

If $\beta = 1$ and $\alpha = 0$, then exchange rate volatility follows the volatility determined solely by economic fundamentals. The exchange rate variance equals the fundamentals variance:

$$\operatorname{var}\left(\Delta_i^e s_{t+i}\right) = \operatorname{var}\left(i_t - i_t^*\right) \qquad (2.3.2)$$

However, if $\alpha \neq 0$ and $\beta < 1$, then there are elements, not included in economic fundamentals, which generate a risk premium. The increase in default risk as a result of high levels of foreign debts affects the risk premium leading it to be more volatile than the exchange rate in terms of what would be expected solely from fundamentals. The expression (2.3.3) reflects that the risk premium variance is greater than the expected exchange rate variance defined by fundamentals.

$$\operatorname{var}(\alpha_t) > \operatorname{var}(\Delta_i^e s_{t+i}) \qquad (2.3.3)$$

A recent, influential study about the relationship between exchange rate volatility and debt stocks is due to Devereux and Lane (2003). They examined a wide set of countries including developing and developed countries using monthly data from 1995:1 to 2000:9. The study followed Mundell's (1961) model of optimal currency areas (OCA) which defines trade interdependence and the degree of commonality in economic shocks as two essential economic preconditions to make regions or countries part of an OCA. Although this paper mostly investigated the effect of optimum currency area variables on exchange rate volatility, it also examined the exchange rate vulnerability to external shocks from large stocks of un-hedged foreign currency denominated debts. A large cross-section model of developed and developing countries was used to identify the main determinants of exchange rate volatility.

The econometric specification was as follows:

$$Vol_{ij}^{ER} = \alpha + \beta X_{ij} + \gamma FIN + \varphi ExtFin_{ij} + \rho(FIN_j \; x \; ExtFin_{ij}) + \varepsilon_{ij} \qquad (2.3.4)$$

where Vol_{ij}^{ER} is the level of bilateral nominal exchange rate volatility between countries i and j, X_{ij} is the set of standard OCA variables, FIN_j is the size of the domestic financial sector and $ExtFin_{ij}$ is a measure of the financial dependence of country j on country i. The volatility of the nominal exchange rate is measured by its standard deviation, the OCA variables are measured by the sum of exports and imports between *i* and *j* countries divided by country *j*'s GDP, and the *ExtFin* variable consists of Bank claims made of own-currency loans.

The results showed that for developed countries the bilateral exchange rate volatility is either positively affected by external financial linkages or affected insignificantly. In contrast, for developing countries financial linkages decrease sharply bilateral exchange rate volatility. The OCA variables, in turn, demonstrated consistent results with the underlying theory so that the greater is bilateral trade the lower is exchange rate volatility.

In short, Devereux and Lane (2003) used a cross section analysis to capture the effect of debts on exchange rate volatility. The cross section modelling is a static analysis which has a limited power to investigate effects that a change in one variable has on another over time and assumes the same data generating process (DGP) for all the cross sectional units without considering country-specific effects. Chapter V proposes to examine the impact of foreign debts on exchange rate volatility using time series modelling. It also investigates if exchange rate volatility is stabilised by monetary policy interventions.

2.4 - Conclusion

This brief overview of theoretical and empirical findings about exchange rate determination leads to the conclusion that even though the methods of analysis have progressed significantly, the results still remain contestable. In fact, a huge number of studies has been produced, but the models continue to be unsatisfactory in giving a robust explanation or to provide unquestionable responses, from using either within-sample or out-of-sample analyses, to the phenomenon of exchange rate movements.

Current investigations about exchange rate behaviour are increasingly focused on improving the performance of econometric approaches. Much progress has been made in developing estimators that essentially provide more accurate parameter estimates based on the traditional monetary model of the exchange rate. An important research field was opened by introducing panel data modelling into investigations. Panel data models offer the advantage of identifying and measuring effects that are not detectable in pure cross-section or time-series models. Furthermore, as panel data models gather more informative data they allow more variability, less collinearity among variables, more degrees of freedom and greater efficiency. These characteristics are particularly relevant to studies on exchange rate behaviour enabling the control of the heterogeneity inherently associated with the economic dynamics of different countries. For instance, a monetary shock may generate heterogeneous impacts on exchange rate movements in different countries. However, aggregation and feedback effects biases emerge when using dynamic panel data models. These difficulties may lead asymptotically to distortions in distributions and affect the test statistics. The estimators need to be improved and therefore there is still much to be done in this area.

The studies related to the presence of bubbles driving the exchange rate from its fundamentals are, in turn, based mostly on the standard monetary model. Although efforts have been made in order to perfect the tests for bubbles, the problem of model misspecifications still persists and may affect the results. Although this problem is an important topic in studies on bubbles, it goes beyond the scope of this thesis. Moreover, the use of stochastic and rational bubble concepts, rather than deterministic, brought considerable advances to the current debate by introducing the bubble process into the information set available to the market. The rationality concept implies basically a data generating process (DGP) for bubbles, which market agents take into account to create a self-fulfilling behaviour diverging from fundamentals. However, investigations are commonly based on linear models and conclusions for the presence of bubbles may be misleading. In fact, the existence of non-linear behaviour in the exchange rate may not be easily detectable by standard econometric tests designed usually for the analysis of linear theoretical models.

In particular, this survey, and the subsequent chapters, pursues the first and the third research groups as defined by MacDonald (1999). The papers discussed analyse the relationship between the exchange rate and fundamentals taking the traditional monetary model as a reference. The two subsequent chapters also study this relationship analysing it in a panel data context with homogeneity relaxed and investigating the presence of bubbles driving exchange rates, using a single equation model based on Markov-switching regimes, respectively. Finally, the last chapter examines the impact of foreign debt shocks on the exchange rate using vector autoregression modelling.

To finish, it is important to emphasise that this survey is just a part of the discussion of exchange rate behaviour. Particularly, it did not try to present the whole debate which involves many more theoretical and empirical questions. The central objective was to introduce essential aspects of relevant studies about exchange rate behaviour that supports the research questions developed in the following chapters.

CHAPTER III

EFFECTS OF THE MONETARY FUNDAMENTALS

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 focussed 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 developed 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, developing 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[15].

Table 3.1 displays some comparative economic figures between developing and developed countries. The figures for the inflation rate and budget balance suggest that developing countries are more subject to monetary shocks which should impact on exchange rates.

Table 3.1 – Some Economic Indicators for Developed and Developing Countries 1975-2000*

Economic Indicators**	Developed Countries	Developing 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.

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

Krugman (1991) points out, on the other hand, that the effect of target zone regimes[16] on exchange rates may be stabilising, even though some fraction of the volatility of relevant market fundamentals may impact on the exchange rate. Thus, monetary shocks might generate smaller effects on the exchange rate under managed regimes. Flood and Rose (1995) reinforce this argument and show that exchange rate regimes based on pure floats are supposed to increase exchange rate volatility more than in managed exchange rate regimes. They demonstrate that in monetary models with flexible prices the conditional volatility of fundamentals is substantially higher during periods of floating exchange rates than during periods of fixed or managed exchange rates. Essentially, this implies that fundamentals volatility is not being transferred onto the exchange rate in managed regimes, and thus it may experience more stable behaviour over time.

This complex debate raises one of the six major puzzles in international macroeconomics posited by Obstfeld and Rogoff (2000b). Basically, they state that no economic model is good enough to give an accurate explanation of exchange rate behaviour.

The previous chapter provided a selective survey of the theory and empirical evidence on exchange rate determination. The goal of this chapter is to give an additional contribution on the ability of the traditional monetary approach in determining nominal exchange rate movements. In essence, the objective is to investigate the existence of long-run relationships between nominal exchange rates and basic components of the monetary model using panel data techniques to offer three relevant contributions.

[16] A target zone regime differs from a fixed rate regime in allowing fairly wide range of variation for the exchange rate around some reference rate (see Williamson, 1985).

The first contribution is this study implements a comparative study between industrialised market economies and emerging market economies using panel data to examine the explanatory power of the traditional monetary approach to exchange rate determination. The basic idea is that industrialised market economies are characterised by robust monetary fundamentals and freely floating exchange rate; whereas emerging market economies present weak monetary fundamentals, but, in general, exchange rate regimes are based on crawling peg (managed). This contrasting economic context may reveal relevant differences and provide an important contribution to the literature.

The analysis is carried out through recently developed procedures in panel data for the detection of unit roots in time series, as well as for identifying the existence of cointegrating relationships between variables. The implementation of both these procedures constitutes the second contribution of this chapter by applying them distinctively to these 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 Lagrange multiplier (LM) and 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.

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 chapter, 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 chapter 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 chapter 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 chapter is structured into five distinct sections. Section 3.1 outlines the main features of the theoretical model underlying the empirical analysis. Section 3.2 describes the statistical data, variables used and sources. Section 3.3 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 3.4 presents the conclusions.

3.1 - 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). The basic difference between the sticky-price and flexible-price models is related to the assumption of continuous fulfilment of the PPP condition. While for the flexible-price assumption, the PPP condition is met continuously, for the sticky-price case this condition is only fulfilled slowly in the long run, generating the well known exchange rate overshooting result. Nevertheless, this is a controversial issue as there are 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 chapter uses three different monetary model versions in which the exchange rate determinants are analysed in flexible-price format[17]. Note that similar theoretical frameworks were used by Frankel (1979), 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).

The strategy of this chapter is to begin from a simple flexible-price monetary model and assume a symmetric treatment between domestic and foreign countries. 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[18]:

$$m_t = p_t + \kappa y_t - \theta i_t \qquad (3.1.1)$$

$$m_t^* = p_t^* + \kappa^* y_t^* - \theta^* i_t^* \qquad (3.1.2)$$

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[17] 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 chapter.

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

equations (3.1.1) and (3.1.2) 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:

$$s_t = p_t - p_t^* \qquad (3.1.3)$$

Given the monetary equilibrium expressed by (3.1.1), (3.1.2) and the PPP condition denoted by (3.1.3), 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:

$$s_t = (m_t - m_t^*) - (\kappa y_t - \kappa^* y_t^*) + (\theta i_t - \theta^* i_t^*) + e \qquad (3.1.4)$$

Equation in (3.1.4) 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 (3.1.4) takes the following form:

$$s_t = \left(m_t - m_t^*\right) - \kappa\left(y_t - y_t^*\right) + \theta\left(i_t - i_t^*\right) + e_t \qquad (3.1.5)$$

Equation (3.1.5) 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 (3.1.5)[19]. 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 (3.1.5) 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

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

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 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 (3.1.5), that is:

$$\Delta^e s_{t+1} = \left(i_t - i_t^*\right) \qquad (3.1.6)$$

where the expected change in the exchange rate ($\Delta^e s$) is proportional to the nominal interest rate differential. Equation (3.1.5) can be suitably rearranged as follows:

$$s_t = (1+\theta)^{-1}(m_t - m_t^*) - \kappa(1+\theta)^{-1}(y_t - y_t^*) + \theta(1+\theta)^{-1} s_{t+1}^e + e_t \qquad (3.1.7)$$

Applying the rational expectations solution to (3.1.7) generates the following representation:

$$s_t = (1+\theta)^{-1}\sum_{i=0}^{\infty}\left(\frac{\theta}{1+\theta}\right)^i E_t\left[(m_{t+i} - m_{t+i}^*) - \kappa(y_{t+i} - y_{t+i}^*)\right] + \varepsilon_t \qquad (3.1.8)$$

where $\varepsilon_t = \phi\left[E_t(s_{t+i} - s_t)\right] + e_t$. If no bubbles are present in the stochastic process then $\varepsilon_t \rightarrow I(0)$. This means that there is a fundamental solution. Equation (3.1.8) constitutes a second version of equation (3.1.5) assuming that the UIRP condition holds.

Finally, from equation (3.1.5) 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 $(\pi - \pi^*)$ in (3.1.5) 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:

$$\Delta^e s = -\varphi\left(s - \bar{s}\right) + \pi - \pi^* \tag{3.1.9}$$

where π and π^* are the current expectations of long-run inflation rates at home and abroad, respectively. Equation (3.1.9) means that expected changes in the exchange rate must be proportional to the current gap in the short run $\left(s - \bar{s}\right)$, and once $s = \bar{s}$ 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 (3.1.9) and rearranging, the following result is obtained:

$$s - \bar{s} = -\varphi^{-1}[\left(i - \pi\right) - (i^* - \pi^*)] \tag{3.1.10}$$

Equation (3.1.10) defines the exchange rate gap as a function of the real interest rate differential. Thus, in the long-run, when $s_t = \bar{s}_t$, then $\bar{i} - \bar{i}^* = \pi - \pi^*$, where $\bar{i}$ and $\bar{i}^*$ denote the long-run interest rates. Given that the purchasing power parity (PPP) condition holds continuously and $(\bar{i} - \bar{i}^*) = (\pi - \pi^*)$, the expression for the long-run exchange rate (3.1.5) can alternatively be defined as follows:

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

$$\bar{s} = \left(\bar{m} - \bar{m}^*\right) - \kappa\left(\bar{y} - \bar{y}^*\right) + \theta\left(\pi - \pi^*\right) + e \tag{3.1.11}$$

If it is also assumed that the equilibrium money supply and income levels are defined by their current levels, equation (3.1.11) can be introduced into (3.1.10) to obtain the following final equation:

$$s = \left(m - m^{*}\right) - \kappa\left(y - y^{*}\right) - \varphi^{-1}\left(i - i^{*}\right) + \left(\varphi^{-1} + \theta\right)\left(\pi - \pi^{*}\right) + e \qquad (3.1.12)$$

From equation (3.1.12) 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[20].

The equations set out in (3.1.5), (3.1.8) and (3.1.12) comprise the economic models investigated in this chapter:

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 $\delta_4 = \left(\varphi^{-1} + \theta\right) = 0$.

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[20] 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.

3.2 - 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).

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 ($\pi-\pi^{*}=\Pi$).

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[21].

3.3 – Econometric Methodology

The procedure of empirical investigation tests the theoretical models as in section 3.1 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;

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

- Estimation of cointegration vectors: The econometric estimation procedure assumes the results found by Phillips and Moon (1999) and uses the pooled 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 3.2.

3.3.1- 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) considered the use of pooled cross-section and time series data to generate powerful unit root tests. They 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 $t = 1,\ldots,T$. 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.

The Levin, Lin and Chu test has 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.

Nevertheless, the Levin, Lin and Chu's specification restricts the parameter δ_i, (the coefficient on the lagged level variable in the unit root test, see Appendix equation A.1.1) to be homogeneous across units. Im, Pesaran and Shin (2003), thus, suggested allowing for heterogeneity related to the value of δ under the alternative hypothesis. This suitable modification aims at capturing the individual realities in panel studies based on a large number of units[22].

The results of both Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003) tests are reported in Tables 3.2 and 3.3 for variables in levels and for variables in first difference, respectively.

Table 3.2 – 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).

[22] 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.

Table 3.3 – 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 3.2 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, is found to be stationary in levels. The LLC test which assumes a homogeneous coefficient for lagged variables in the ADF regression, and the IPS test which, in turn, assumes a heterogeneous coefficient for the lagged variables, reveal that the exchange rate series is stationary in levels for industrialised market economies. 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[23].

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 3.1) do not hold for both sets of countries. Finally, for the 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 3.3 reports the results obtained from the LLC and IPS tests at the 5% and 1% significance levels[24].

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 (A.1.8 in the Appendix)

[23] Note, nonetheless, that it is used a fixed lag order p_i=4 in the ADF regression A.1.8 (see in the Appendix). The tests were also applied by using lag order p = 1, 2, 3, but no significant difference was detected.
[24] Note also that the inclusion of a trend did not change this conclusion.

might lead to misleading conclusions[25]. 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[26]. Next, the IPS formula (A.1.9 in the Appendix) is employed for different values of $E\{t_{iT}(p_i) \mid \delta_i = 0\}$ and $Var\{t_{iT}(p_i) \mid \delta_i = 0\}$ tabulated by Im *et al.* (2003) for different *Ts* and lag orders. Table 3.4 displays the lag order used country to country.

Table 3.4 – 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

[25] 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.

[26] 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[27]. 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:

$$\Delta y_{it} = \alpha_i + b_i y_{i,t-1} + c_i \bar{y}_{t-1} + \sum_{j=0}^{p} d_{ij} \Delta \bar{y}_{t-j} + \sum_{j=1}^{p} \Delta y_{i,t-j} + \varepsilon_{it} \qquad (3.3.1.1)$$

where $\bar{y}_{t-1}$ and $\Delta\bar{y}_{t-j}$ in (3.3.1.1) 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 3.4 generated $CIPS_1$ = -1.788 and $CIPS_2$ =

[27] 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.

-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[28].

3.3.2 - 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 research has taken two distinct directions. The first one comes from an analogue panel model linked to the ideas of the pioneering work of Engle and Granger (1987)[29]. It consists of using a static regression model for constructing residuals-based test statistics and tabulating distributions. Accordingly, this first approach assumes the null hypothesis of no cointegration, and it is linked to Pedroni's papers (see Pedroni, 1995, 1997) and more recently to Kao (1999). The second one is developed by McCoskey and Kao (1998) which also constructs a residual-based test statistic, but as opposed to the first one, assumes the null hypothesis of cointegration.

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

[28] 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.

[29] 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.

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 chapter, 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)[30].

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

[30] Note that, all statistics are constructed using residuals generated by the regression A.2.1 (see in the Appendix) and by the use of nuisance parameters estimators.

hypothesis and diverges to negative infinity 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 3.1. 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 3.5).

Table 3.5 – 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 : $er = (m2\text{-}m2^*) - (y - y^*) + e$**
[2] **Model II : $er = (m2\text{-}m2^*) - (y - y^*) + (ir - ir^*) + e$**
[3] **Model III : $er = (m2\text{-}m2^*) - (y - y^*) + (ir - ir^*) + (\pi - \pi^*) + e$**

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)[31] 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 3.6 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 3.6).

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

Table 3.6 – 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 : $er = (m2\text{-}m2^*) - (y - y^*) + e$**
[2] **Model II : $er = (m2\text{-}m2^*) - (y - y^*) + (ir - ir^*) + e$**
[3] **Model III : $er = (m2\text{-}m2^*) - (y - y^*) + (ir - ir^*) + (\pi - \pi^*) + e$**

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 3.7 and 3.8 display the results obtained for industrialised economies and emerging economies, respectively[32]. 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 3.7 and 3.8).

Table 3.7 – 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**

[32] 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 3.8 – 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*) + ($\pi - \pi$ *) + e**

3.3.3 – Long-run Average Relations in Non-stationary Panel Data

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 chapter. Phillips and Moon (1999) demonstrated that long-run average relations between integrated panel vectors can be found even if no cointegration is detected. According to them, a limit theory allows for the testing of hypotheses about the long-run average parameters both within and between subgroups of the entire population. They investigated regressions with nonstationary panel data for which the time series component is an integrated process and where the *T* and *N* dimensions are large. The underlying idea is based on the fact that panel data can distinguish effects that time series or cross section data cannot identify individually.

Suppose, for example, the existence of two I(1) random vectors, $Y_{i,t}$ and $X_{i,t}$, which are assumed not to have a cointegrating relationship, and a time series regression for a given unit i is carried out. This regression for i will produce a regression coefficient that has a non-degenerate limit distribution characterising a result that the econometric literature usually considers as potentially spurious [see Granger and Newbold (1974) and Phillips (1986)]. Nevertheless, if the panel regression of $Y_{i,t}$ on $X_{i,t}$ with large cross sectional and time series dimensions is estimated the noise in the time series data can often be assumed as independent across individuals, even if the noise is strong. Phillips and Moon state that this result is reached by pooling the cross section and time series observations which can reduce the strong effect of the residuals in the regression while retaining the strength of the vector $X_{i,t}$. Hence, it is expected that a panel-pooled regression provides consistent estimates of the long-run regression coefficients.

The Phillips and Moon's study is closely related to work developed by Pesaran and Smith (1995) who investigated the impact of nonstationary variables on cross section regression estimates. Pesaran and Smith concluded that provided the regressors are exogenous, and the disturbances are independent, no spurious correlation arises by running a cross section regression with a finite number of time series observations. Phillips and Moon extended Pesaran and Smith's investigation to a more general context when $T \to \infty$ and $N \to \infty$ sequentially or jointly in panel regressions, and thus derived the long-run average relationship in terms of a matrix regression coefficient from the cross section long-run average covariance matrix[33].

The relevance of this topic for investigating the relationship between the exchange rate and monetary fundamentals is related to panel unit root and panel cointegration tests. The cointegration tests applied in this chapter supported results in favour of the null hypothesis of no cointegration. It leads straightforwardly to the conclusion that monetary fundamentals are not able to explain exchange rate movements in the long. However, according to Phillips and Moon even if no cointegration relation is found the use of a panel estimation procedure may generate consistent long-run coefficients. The following section shows the results obtained by using the pooled mean group estimation procedure for the exchange rate and its determinants using Phillips and Moon in support of this procedure.

3.3.4 - Method of Econometric Estimation

The pooling data technique is a methodology whereby 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

[33] The demonstration presented by Phillips and Moon on how two component random vectors $Y_{i,t}$ and $X_{i,t}$ of $Z_{i,t}$ may have a long-run average relation even if no cointegration relation is defined is somewhat complex and is not presented in this chapter.

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 (see Appendix).

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 3.3.4.1 presents the main empirical findings from applying the PMG procedure to the exchange rate and monetary fundamentals.

3.3.4.1 – PMGE and Exchange Rate Fundamentals

The pooled mean group (PMG) estimation procedure is applied to the theoretical model discussed in Section 3.1. 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[34]. Finally, an empirical autoregressive distributed lag (ARDL) equation is constructed for the three specifications presented in the Section 3.1. The full long-run exchange rate specification is given by:

$$er_{it} = \theta_{0i} + \theta_{1i}\left(m_{it} - m_{it}^{*}\right) - \theta_{2i}\left(y_{it} - y_{it}^{*}\right) + \theta_{3i}\left(ir_{it} - ir_{it}^{*}\right) + \theta_{4i}\left(\pi_{it} - \pi_{it}^{*}\right) + u_{it} \qquad (3.3.4.1)$$

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

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_{1i} is expected to be one if

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

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:

$$er_{it} = \mu_i + \lambda_{1iq}\sum_{q=1}^{3} er_{it-q} + \delta_{2iq}\sum_{q=0}^{3} m_{it-q} - \delta_{3iq}\sum_{q=0}^{3} y_{it-q} + \delta_{4iq}\sum_{q=0}^{3} ir_{it-q} + \delta_{5iq}\sum_{q=0}^{3} \pi_{it-q} + \varepsilon_{it} \qquad (3.3.4.2)$$

and the general error correction representation - ECM of (3.3.4.2) is given by:

$$\Delta er_{it} = \phi_i\left(er_{it-1} - \theta_{0i} - \theta_{1i}\, m_{it-1} + \theta_{2i}\, y_{it-1} - \theta_{3i}\, ir_{it-1} - \theta_{4i}\, \pi_{it-1}\right) + \delta_{1iq}\sum_{q=1}^{3} \Delta er_{it-q} + \delta_{2iq}\sum_{q=0}^{3} \Delta m_{it-q}$$
$$-\delta_{3qi}\sum_{q=0}^{3} \Delta y_{it-q} + \delta_{4qi}\sum_{q=0}^{3} \Delta ir_{it-q} + \delta_{5qi}\sum_{q=0}^{3} \Delta \pi_{it-q} + \varepsilon_{it} \qquad (3.3.4.3)$$

where from (3.3.4.2):

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

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

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

From equation 3.3.4.3 three different error correction empirical models can be derived, as defined theoretically in Section 3.1. 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 developed 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[35].

The econometric results were organised in three different groups: the first one reports country by country long-run estimates

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

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 3.3.4.3) 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 3.9). 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 3.9). The long-run coefficients, estimated by MG and by PMG, were positive for M1 and negative for real income for the summations of countries as whole and this result met the expected signs according to the theoretical economic model I (see Table 3.10).

Table 3.9 – 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 3.10 – 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 3.9). 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 3.10). 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 3.10)[36]. 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 3.9). 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 3.10). 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 3.10). 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 3.1).

[36] According to the theoretical literature (see Section 3.1) 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 3.10).

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 3.10). 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[37].

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

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

reported the coefficient signs were not theoretically consistent (see Table 3.11). 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 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 3.11). 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 3.11). 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 3.12). 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 3.12). 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 3.11 – 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 3.12 – 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 3.12). 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 $\bar{R}^2$ 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 3.13 and 3.14 present these computations for industrialised and emerging market economies, respectively[38].

Table 3.13 – 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

[38] The investigation developed in this chapter 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.

Table 3.14 – R^2 and $\overline{R}^2$ Estimated by PMG Using M1 - Emerging Economies

Countries	Model I		Model II		Model III	
	R^2	$\overline{R}^2$	R^2	$\overline{R}^2$	R^2	$\overline{R}^2$
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.

3.4 - Conclusion

The central aim of this chapter 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.

Three different versions of monetary models for exchange rate determination were investigated: The first, a basic version, consisted of money supply and real income variables. Two extensions of this model added the interest rate and expected future inflation rate differentials. Additionally, using data in panel form, two different groups of countries were investigated: ten industrialised market economies and seven emerging market economies. The objective was to investigate the ability of the monetary approach to explain exchange rate movements in different economic environments. The macroeconomic reality between these two sets of countries differs significantly in terms of inflation, budget balance, etc. Thus, it was expected that the monetary approach to the exchange rate would have different explanatory power for these distinct economies.

Most of the panel data unit root tests confirmed that the variables used in this investigation were I(1) at the 5% and 1% significance level. The exchange rate and interest rate of industrialised market economies were found to be stationary in levels so that the UIRP condition holds. For panel cointegration tests, the results 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 tests.

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. The results for industrialised market economies were reinforced as the coefficients estimated by PMG for the money supply were in accordance with the theoretical economic model. For emerging market economies, the results were weaker as only Chile demonstrated a consistent long-run relationship between the exchange rate and monetary fundamentals.

The positive sign found on the interest rate coefficient was in favour of a flexible-price monetary model and this prevailed for both sets of economies. This is an important result as it implies a continuous PPP. Furthermore, the statistical insignificance of the expected inflation rate coefficient led to the conclusion that this variable is not relevant in explaining long-run exchange rate movements. This finding reinforces the hypothesis that inflation expectations are already included in the long-run fundamental equation of the exchange rate.

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.

In summary, the discussion about the relation between exchange rate behaviour and monetary fundamentals remains open. Even though the results were somewhat mixed, this chapter was able to find some support for the monetary model of exchange rate determination However, the estimated lower coefficient values than the theory predicted suggests that the monetary approach is only able to give a partial explanation to exchange rate volatilities and more robust results are still needed. The debate about exchange rate determination has mach 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, and other market frictions. Chapter IV of this thesis moves on to investigate the presence of bubbles in the foreign exchange market using the Markov-switching regime approach.

APPENDIX

A.1 - Panel Unit Root Test

Levin, Lin and Chu (2002)

Assuming that the stochastic process y_{it} is generated by the following models:

$$\Delta y_{it} = \alpha_{0i} + \theta_t + \alpha_{1i}t + \delta_i y_{it-1} + \zeta_{it} \qquad \text{(A.1.1)}$$

where $-2 < \delta_i \leq 0 \ \forall i = 1,...,N$ and if $\delta_i = 0$, then $\alpha_{1i} = 0$. The disturbance ζ_{it} is distributed independently across units and follows a stationary invertible ARMA process for each unit.

The traditional ADF test for data panel forms uses t-statistic to test the null hypothesis that $\delta_i = 0 \ \forall i = 1,...,N$. In particular, the series y_{it} may include an individual-specific mean and time trend and the test procedure consists of evaluating the null hypothesis that $\delta_i = 0$ in (A.1.1) and $\alpha_{1i} = 0 \ \forall i = 1,...N$ against the alternative hypothesis that $\delta_i < 0$ and $\alpha_{1i} \in \mathbb{R} \ \forall i = 1,...,N$[39].

The strategy adopted by LLC, in turn, consists of computing a panel test statistic in which under the null hypothesis

[39] The researcher should be aware to the presence or not of a deterministic element, for example, an intercept or time trend, in the data. If this element is present in the data but it is not included into the regression procedure, then the unit root test will be inconsistent (see Campbell and Perron, 1991).

the normalised residual innovations $\tilde{e}_{it}$[40], $\tilde{e}_{it} = \frac{\hat{e}_{it}}{\hat{\sigma}_{e_i}}$, are independent of the normalised lagged residuals $\tilde{v}_{it-1}$[41], $\tilde{v}_{it} = \frac{\hat{v}_{it}}{\hat{\sigma}_{e_i}}$, for each unit in the panel [see Levin, Lin and Chu (2002) for details]. Thus, the following regression can be performed to test the null hypothesis:

$$\tilde{e}_{it} = \delta \tilde{v}_{it-1} + \tilde{\varepsilon}_{it} \qquad \text{(A.1.2)}$$

where $N\tilde{T}$[42] corresponds the total number of observations in the regression.

By performing the regression (A.1.2) the following statistic can be derived:

$t_{\delta} = \frac{\hat{\delta}}{RSE\left(\hat{\delta}\right)}$, the regression t-statistic for the null hypothesis[43]

(A.1.3)

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40 $\hat{e}_{it} = \Delta y_{it} - \sum_{L=1}^{p_i} \hat{\theta}_{1iL} \Delta y_{it-L} - \alpha_{1i} d_t$

41 $\hat{v}_{it-1} = y_{it-1} - \sum_{L=1}^{p_i} \hat{\theta}_{1iL} \Delta y_{it-L} - \hat{\alpha}_{1i} d_t$

42 $\tilde{T} \equiv \left(T - \overline{p} - 1\right)$ is the average number of observations per group in the panel and $\overline{p} \equiv N^{-1} \sum_{i=1}^{N} p_i$ is the average lag order for ADF regressions per group.

According to Levin, Lin and Chu, as some of the deterministic elements can be present in the data, so the limiting distribution for the statistic (A.1.3) can diverge to negative infinity. In order to correct such a distortion the following adjusting test statistic can be used:

$$t_{\delta}^{*} = \frac{t_{\delta} - N\tilde{T}\hat{S}_{NT}\hat{\sigma}_{\varepsilon}^{-2}RSE\left(\hat{\delta}\right)\mu^{*}\tilde{T}}{\sigma_{\tilde{T}}^{*}} \qquad \text{(A.1.4)}$$

where $\mu_{\tilde{T}}^{*}$ and $\sigma_{\tilde{T}}^{*}$ are the mean adjustment and standard deviation adjustment for a given deterministic specification and time series dimension $\tilde{T}$. **These adjustments allow the test statistic t_{δ}^{*} to have a $N(0,1)$ distribution.**

Im *et al.* (2003)

Assuming the same representation for the stochastic process as in (A.1):

$$\Delta y_{it} = \alpha_{0i} + \alpha_{1i}t + \delta_i y_{it-1} + \zeta_{it} \quad i = 1,\dots,N; \quad t = 1,\dots,T \qquad \text{(A.5)}$$

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[43] $\hat{\delta} = \dfrac{\sum\limits_{i=1}^{N}\sum\limits_{t=2+p_i}^{T}\tilde{v}_{it-1}\tilde{e}_{it}}{\sum\limits_{i=1}^{N}\sum\limits_{t=2+p_i}^{T}\tilde{v}_{it-1}^{2}}$, $\tilde{\sigma}_{\varepsilon} = \left[N\tilde{T}^{-1}\sum\limits_{i=1}^{N}\sum\limits_{t=2+p_i}^{T}\left(\tilde{e}_{it} - \hat{\delta}\tilde{v}_{it-1}\right)^{2}\right]^{1/2}$, and

$RSE\left(\hat{\delta}\right) = \hat{\sigma}_{\varepsilon}\left[\sum\limits_{i=1}^{N}\sum\limits_{t=2+p_i}^{T}\left(\tilde{v}_{it-1}\right)^{2}\right]^{-1/2}$.

The null and alternative hypotheses are defined as:

$$H_0: \delta_i = 0 \quad \forall i \tag{A.1.6}$$

against the alternative:

$$H_A: \delta_i < 0,\ i = 1,\dots,N_1;\quad \delta_i = 0,\ i = N_1+1, N_1+2,\dots,N \tag{A.1.7}$$

Hence, in order to test for the null hypothesis in (A.1.6), Im *et al.* (2003) proposed a group-mean Lagrange multiplier (LM) statistic and a group-mean t-statistic test. The group-mean t-statistic test is used in this chapter. Again starting point is the ADF regression:

$$\Delta y_{it} = \delta_i y_{it-1} + \sum_{L=1}^{p_i} \theta_{ij} \Delta y_{it-L} + \alpha_i + \varepsilon_{it} \quad t = 1,\dots,T \tag{A.1.8}$$

Thus, Im *et al.* (2003) propose the following statistic which is known as group-mean t-statistic test:

$$\Psi_{\bar{t}} = \frac{\sqrt{N}\left\{\bar{t}_{N,T} - N^{-1}\sum_{i=1}^{N} E\left[t_{i,T}(p_i,0) \mid \delta_i = 0\right]\right\}}{\sqrt{N^{-1}\sum_{i=1}^{N} Var\left[t_{i,T}(p_i,0) \mid \delta_i = 0\right]}} \tag{A.1.9}$$

where $\bar{t}_{N,T}$ is the average of the N cross-section ADF(p_i) t-statistic, $E\{t_{iT}(p_i) \mid \delta_i = 0\}$ and $Var\{t_{iT}(p_i) \mid \delta_i = 0\}$ are the mean and the variance of the average ADF(p_i) statistic under the null hypothesis, respectively. Monte Carlo simulations show under that $H_0: \delta_i = 0,\ \forall i$,

$\Psi_{\bar{T}} \Rightarrow N(0,1)$, that is, converge to a normal limiting distribution as $T, N \rightarrow \infty$ and $N/T \rightarrow k$ (the rate of convergence) where k is a finite positive constant.

A.2 - Panel Cointegration Test

This method consists of using the residuals from the cointegrating regression based on the following representation:

$$Y_{i,t} = \alpha_i + \delta_i t + \beta_{Mi} X'_{Mi,t} + e_{i,t} \qquad \text{(A.2.1)}$$

$$i = 1,2,...,N \;\; ; \;\; t = 1,2,...,T \quad and \quad m = 1,2...,M$$

where $\beta_i = \left(\beta_{1i}, \beta_{2i}, ..., \beta_{mi}\right)$, $Y_{it} = \left(y_{1it}, y_{2it}, ..., y_{mit}\right)$, $X_{it} = \left(x_{1it}, x_{2it}, ..., x_{mit}\right)'$. T refers to the number of observations over time, N denotes the number of cross-sections in panel, and M refers to the number of regression variables. It is worth emphasising that such a formulation is flexible in the sense that the underlying heterogeneity in panel includes heterogeneity in slope coefficients $\left(\beta_{1i}, \beta_{2i}, ..., \beta_{mi}\right)$ as well as fixed effects (α_i) and individual specific deterministic trends ($\delta_i t$)[44].

[44] According to Pedroni (1999) the inclusion of individual-specific fixed effects or individual-specific time trends causes some difficulties by affecting the asymptotic distribution and respective critical values. In case of unit root presence, for instance, specific-fixed effects and specific-time trends introduced in the regression cause the sample average over time to diverge from the population means at the rate $\sqrt{T}$ as T grows large.

Pedroni takes the small sample performances to derive the asymptotic distributions of seven different statistics, that is, four referred to as the within-dimension, and three referred to as the between-dimension. Conceptually, the construction of within-dimension statistics is based on the sum of both the numerator and the denominator terms over the N dimensions separately. For the construction of between-dimension statistics, in turn, the procedure consists of first dividing the numerator by the denominator prior to summing over N dimensions. Consequently, the within-dimension statistics are based on estimators that pool the autoregressive coefficient across different units in the panel, whereas the between-dimension statistics have as a base estimators which average the individually estimated coefficients for each unit i. In essence, the tests are based on two distinct error processes from (A.2.1) as the following equations:

$$\hat{e}_{i,t}=\hat{\gamma}_i\hat{e}_{i,t-1}+\hat{u}_{i,t} \quad \textit{and} \quad \hat{e}_{i,t}=\hat{\gamma}_i\hat{e}_{i,t-1}+\sum_{k=1}^{k_i}\hat{\gamma}_{i,k}\Delta\hat{e}_{i,t-k}+\hat{u}^*_{i,t} \qquad \text{(A.2.2)}$$

This distinction has two important meanings related to the autoregressive coefficient of the estimated residuals under the alternative hypothesis of cointegration. Firstly, the within-dimension statistics for analysis of cointegrating vectors test the null hypothesis $H_0: \gamma_i = 1 \ \forall i$ in (A.2.2) against the alternative hypothesis $H_1: \gamma_i = \gamma < 1 \ \forall i$, so that it is presumed a common value for $\gamma_i = \gamma$ across different units in the panel. Secondly, the between-dimension statistics, by contrast, assume the null hypothesis of a residual-based test as $H_0: \gamma_i = 1 \ \forall i$ versus the alternative $H_1: \gamma_i < 1 \ \forall i$, that is, as opposed to the former one these statistics allow for an additional heterogeneity since it is not presumed a common value for $\gamma_i = \gamma$ under alternative hypothesis. Pedroni himself calls the within-dimension statistics as panel cointegration statistics, and the between-dimension statistics as group mean panel cointegration statistics.

A.3 - Pooled Mean Group Estimation

The basic framework of the pooled mean group estimator is briefly summarised as follows. Supposing, initially, that the overall data set is composed of a time series and cross section data set on time period $t = 1, 2, ..., T$ and units $i = 1, 2, ..., N$, respectively, then the objective is to estimate the following ARDL $(p,q,...,q)$ model:

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{it-j} + \sum_{j=0}^{q} \delta'_{ij} x_{it-j} + \mu_i + \varepsilon_{it} \qquad \text{(A.3.1)}$$

where x_{it} $(k\,x\,1)$ is the vector of explanatory variables for units i; μ_i denote the fixed effects; the coefficient of the lagged dependent variables, λ_{it} are scalars; and δ_{it} are $k\,x\,1$ coefficient vectors. T dimension must be large enough so that one can estimate the model for each unit separately[45]. Accordingly, PSS suggest that the following reparameterisation in (A.3.1) can be carried out generating the representation as follows:

[45] The use a common T and p across units, a common q across units and regressors is not a necessary requirement. The PMG estimator allows the use of different T, p, q across units without affecting the statistical properties.

$$\Delta y_{it} = \phi_i y_{it-1} + \beta_i^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta x_{it-j} + \mu_i + \varepsilon_{it} \qquad (A.3.2)$$

where $\phi_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right)$, $\beta_i = \sum_{j=0}^{q} \delta_{ij}$ and:

$$\lambda_{ij}^{*} = - \sum_{m=j+1}^{p} \lambda_{im}, \qquad j = 1, 2, ..., p-1$$

$$(A.3.3)$$

$$\delta_{ij}^{*} = - \sum_{m=j+1}^{q} \delta_{im}, \qquad j = 1, 2, ..., q-1$$

The time series data can still be stacked for each unit in order to have a stacked representation as follows:

$$\Delta Y_{it} = \phi_i Y_{i,-1} + \beta_i X_i + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta Y_{i,-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta X_{i,-j} + \mu_i U + \varepsilon_i \qquad (A.3.4)$$

where $Y_i = (y_{i1}, ..., y_{iT})^{'}$ is a $T x 1$ vector of the observations on the dependent variables of the ith unit, $X_i = (x_{i1}, ..., x_{iT})^{'}$, is the a $T x k$ matrix of observations on the regressors that vary across units and time periods, $U = (1, ..., 1)^{'}$ is a $T x 1$ vector of 1s, $Y_{i,-j}$ and $X_{i,-j}$ are j period lagged values of Y_i and X_i, and $\Delta Y_i = Y_i - Y_{i,-1}$, $\Delta X_i = X_i - X_{i,-1}$, $\Delta Y_{i,-j}$ and $\Delta X_{i,-j}$ are j period lagged values of ΔY_i and ΔX_i, respectively, and $\varepsilon_i = (\varepsilon_{i1}, ..., \varepsilon_{iT})^{'}$.

It is also assumed that the disturbances ε_{iT} are distributed independently across units and t, with means zero and variance $\sigma^2 > 0$, fourth-order moments, and distributed independently of the regressors, X_{it}. This assumption is equivalent to stating that $\phi_i < 0$ and therefore there exists a long-run relationship between y_{it} and x_{it}.

As the assumption of existence of long-run relationships is one of the major hypotheses of this approach, the long-run coefficient on X_i, $\theta_i = -\beta_i/\phi_i$, are the same across units, that is:

$$\theta_i = \theta, \qquad i = 1, 2, \ldots, N.$$

The approach used to estimate is a likelihood approach which it is assumed that the disturbances ε_{it} are normally distributed. This allows the likelihood of the panel data model being written as products of the likelihoods for each unit. Then a log-likelihood function can be formulated which has the following representation:

$$L_T(\varphi) = -\frac{T}{2}\sum_{i=1}^{N}\ln 2\pi\sigma_i^2 - \frac{1}{2}\sum_{i=1}^{N}\sigma_i^{-2}\left(\Delta Y_i - \phi_i\xi(\theta)\right)' H_i\left(\Delta Y_i - \phi_i\xi(\theta)\right) \qquad \text{(A.3.5)}$$

where:

$$H_i = I_T - W_i\left(W_i' W_i\right)^{-1} W_i',\ \varphi = \left(\theta', \phi', \sigma'\right)',\ \phi = \left(\phi_1, \phi_2, \ldots, \phi_N\right)' \ and\ \sigma = \left(\sigma_1^2, \sigma_2^2, \ldots, \sigma_N^2\right)'.$$

For estimation purpose, the long-run coefficients, θ, and the unit specific coefficients, ϕ_i, can be computed by maximizing the log-likelihood function in (A.3.5) with respect to φ. Hence, the maxi-likelihood estimators (MLE) used to compute the parameters of the model are termed the pooled mean group (PMG) estimators to characterise both the pooling due to the assumption of homogeneity on the long-run coefficients, and the averaging process across units designed to compute means of the estimated error-correction coefficients and the other short-run parameters included in the model.

CHAPTER IV

EFFECTS OF RATIONAL SPECULATIVE BUBBLES

Introduction

In the previous chapter, panel data tests applied to industrialised and emerging market economy countries demonstrated that exchange rate movements were driven only partially by economic fundamentals. For industrialised or emerging market economy countries panel cointegration tests shown that monetary fundamentals were not able to give a convincing explanation for the long-run behaviour of the exchange rate. For industrialised market economy countries the results were found better than for emerging countries. Although the estimates for industrialised countries met, in general, the theoretical predictions for the long-run coefficients, there was a very low correlation between the nominal exchange rate and monetary fundamentals. For emerging market economy countries the results were poor and not consistent with the predicted values.

Since the tests have rejected the existence of robust long-run relationships between the exchange rate and monetary fundamentals, an alternative possibility arises. This consists of investigating whether the presence of bubbles drives the stochastic processes underlying theses variables. This task will be the central theme of this chapter.

A rational speculative bubble in the exchange rate is characterised by an explosive path. Explosive behaviour leads the exchange rate to diverge from the equilibrium level defined by monetary fundamentals. The underlying idea about the presence of bubbles concerns the fact that speculators and investors maintain a belief that, despite the currency being overvalued with respect to its fundamentals, it is still profitable to buy additional units of it. Thus, a bubble acquires importance in driving the exchange rate away from the equilibrium determined by its fundamentals because market agents perceive the presence of profit opportunity. In fact, movements of asset prices away from their fundamentals signal the occurrence of self-fulfilling prophecies of market participants, caused by events that are exogenous to the market. Originally, the possibility of bubbles is credited by some researchers as a contribution of Keynes' (1936) description of equity markets in a context whereby speculators anticipate what average opinion will be rather than a fully informed view of the market. This typical market behaviour leaves the fundamentals information to one side. Essentially, the presence of bubbles in asset markets implies that market participants are not allocating their savings to the best possible investment. Furthermore, the analysis of rational bubbles, based on rational expectations, has an element of indeterminacy, which usually arises when the current decisions of agents depend both on the current market price and on their expectations of future prices (see Obstfeld and Rogoff, 1997). Given this, an obvious conclusion is that a single hypothesis cannot encompass sequences of prices as only one sequence is the market fundamental price path. The other sequences maybe price bubble sequences (see Blanchard and Fischer, 2001).

In empirical research the traditional literature has followed three different methods of investigating the idea of bubbles driving asset prices away from their economic fundamentals equilibrium. The first one is the variance bound test, also known as volatility test, and it was originally proposed by Shiller (1979) to study the volatility of long-term bonds. Shiller's method was employed by Huang (1981) to examine the presence of bubbles in the Deutsch mark exchange rate. The volatility test is based on the idea that if prices have a bubble component, causing them to deviate from the fundamentals solution, then they will vary even if fundamentals are unchanged, or do not move much. In this case, too much price variation suggests the presence of a bubble. Therefore, this class of test is essentially a test of excess volatility in asset prices. The excess volatility of the actual exchange rate relative to the volatility of the exchange rate based on the fundamentals solution is suggestive of the presence of a speculative bubble (see Blanchard and Fischer, 2001).

A second econometric method is based on the Hausman (1978) test and called the Hausman specification test. The specification test was originally published in 1987 by Kenneth West to investigate the presence of bubbles in the stock market. His method was applied to the exchange rate analysis by Meese (1986), and embraces a technique proposed by McCallum (1976). McCallum's technique relies on the use of an instrumental variable, where the unobserved expectation of the exchange rate $E[s_{t+1}/\theta_t]$, based on the information set θ at time t, is replaced by its actual value s_{t+1} minus a forecast error η_{t+1} uncorrelated with θ_t. Essentially the null hypothesis of no bubbles is tested by comparing two different formulations of a forward-looking exchange rate model, where only one of them is consistent with the presence of bubbles.

The major difficulty of employing either the volatility test or the specification test is related to the fact that both methodologies depend strongly on the chosen model of exchange rate determination. Hence, the excess volatility observed in some studies may be caused by factors other than the presence of rational bubbles. In particular for the specification test, Flood and Hodrick (1990) argue that the reliability of conclusions based on omitted variables may be weak since the monetary model in general produces misleading results. This may lead bubble tests towards rejection of the null hypothesis of no bubbles. Flood, Hodrick and Kaplan (1987) find substantive evidence of misspecification of the model used by West which leads to misleading conclusions. Furthermore, according to Flood and Hodrick (1990), bubble tests of the equity market require a well-specified model of equilibrium expected returns that has yet to be discovered. Hence, similar to the excess volatility tests, if the premise of false models holds, then rejection of the null hypothesis of no bubbles by specification tests cannot be attributed solely to bubbles.

The presence of rational bubbles driving the exchange rate away from economic fundamentals can also be analysed by the use of unit root and cointegration tests. This methodology consists of examining the possibility of cointegration between the exchange rate and fundamentals (see MacDonald and Taylor, 1993). In essence, the argument is that if asset prices are not more explosive than their determinants, then it may be inferred that rational bubbles are not present as they would generate an explosive component to asset prices. Thus the strategy consists of implementing unit root tests and cointegration tests to analyse the presence of explosive rational bubbles. These types of consideration motivated the test for bubbles carried out by Diba and Grossman (1988) and MacDonald and Taylor (1993), inter alia.

Bubbles tests based on unit root tests and cointegration tests have, nevertheless, been potentially misleading in the presence of bubbles collapsing periodically. Evans (1991) was the first to acknowledge this kind of difficulty in examining bubbles using either unit root tests or cointegration tests. According to him the traditional cointegration tests leads to rejection of the null hypothesis of no bubbles more often than if the presence of a particular type of bubble was not taken into account. The central argument described in his paper is that unit root tests or cointegration tests, when faced with periodically collapsing rational bubbles may lead, with a high degree of probability, to a wrong conclusion that bubble components are nonexistent in the stochastic process.

Although the arguments developed by Evans have been related to stock price behaviour it can also be extended to exchange rate movements. Sarno and Taylor (2002) state that the concept of periodically collapsing bubbles is important in analysing the behaviour of exchange rates since they collapse almost in finite time. They argue that the essence of the problem is that the presence of periodically collapsing bubbles causes a greater degree of skewness and excess kurtosis in the exchange rate series relative to the economic fundamentals series. As an alternative strategy, Sarno and Taylor suggested making use of a test for non-stationarity which allows the collapse in bubbles to be attributed to sudden movements in non-normal error terms rather than to estimated coefficients of the autoregressive model. Taking up this idea, Taylor and Peel (1998) examined the performance of a new cointegration test in the presence of periodically collapsing bubbles. Such a test statistic gives considerable weight to the skewness and excess kurtosis injected into the data as a consequence of the presence of a bubble.

Essentially, Taylor and Peel proposed a test based on a modification to the least squares estimator in order to make it robust to the presence of error terms exhibiting strong skewness and kurtosis. The mechanics consists of estimating a regression equation by ordinary least squares (OLS) augmented by a covariate vector which captures the effects of skewness and excess kurtosis. This procedure of estimation was proposed, initially, by Im (1996) and denoted the residual-augmented least squares (RALS) estimation procedure (see Section 2.2, Chapter I). A similar test procedure was adopted by Sarno and Taylor (1999) to investigate the presence of periodically collapsing bubbles in the East Asian stock market. Sarno and Taylor tested for stationarity of the log dividend-price ratio and the *ex-post* rate of return, and then tested cointegration for both series. The underlying idea was that if stationarity or cointegration were found between the series the hypothesis of stock price bubbles could be rejected. The results reached by Sarno and Taylor did not reject the hypothesis of bubbles in the East Asian stock market.

More recently, Psaradakis *et al.*(2001) used a new procedure for detecting the presence of periodically collapsing bubbles based on random-coefficient autoregressive models. In fact, they proposed a simpler test procedure which is based essentially on the class of nonstationary varying-coefficient autoregressive models with a stochastic unit root (see Section 2.2, Chapter I)[46].

Another test procedure for periodically collapsing bubbles was carried out by Hall *et al.* (1999) which allowed for the possibility of changes in the dynamic behaviour of asset prices across the sample. In effect, the methodology proposed constitutes a generalisation of the augmented Dickey-Fuller (ADF) unit root test applied to the class of Markov-switching regime models (see Hamilton 1988, 1989). An important difficulty is to identify collapsing periods from expanding ones. Hall *et al.* argue that the proposed methodology overcomes some of the econometric problems involved traditional unit root tests (see Evans 1991). It allows the ADF regression coefficients to switch values between different regimes/states as a consequence of the dynamics of a periodically collapsing bubble. They applied the ADF-switching unit root test to investigate the presence of an explosive autoregressive root to the

[46] This class of model was also examined by McCabe and Tremayne (1995), Leybourne et al. (1996), Granger and Swanson (1997).

hyperinflation process in Argentina during the 1980s (see Section 2.2, Chapter I).

This chapter embraces the approach based on the unit root and cointegration tests, but uses the Markov-switching regime methodology in order to allow for a more robust analysis of periodically collapsing bubbles. Moreover, taking as a starting point the results reached in the previous chapter, the contribution of this chapter is to investigate the hypothesis of a periodically collapsing bubble underlying the movement of the exchange rate. As the econometric results of the previous chapter only provide partial evidence in favour of monetary fundamentals for the determination exchange rates, a set of four industrialised market economy countries (Canada, France, Germany and the United Kingdom) are selected as a sample to examine the presence of speculative rational bubbles[47]. In doing so, the ADF-switching unit root test proposed by Hall *et al.* (1999) is appropriate since this methodology allows consideration of different regimes/states typical of periodically collapsing bubbles in unstable economies. Also, the unit root test approach offers an additional and convenient way of constructing tests of hypotheses of unit roots against alternatives in the sense that the number and the location of change points are unknown. Such pathology is ideally suited to the adoption of the Markov-Switching (MS) methodology.

Although the MS unit root test is an advanced tool to detect the presence of rational bubbles in the exchange rate and in its economic fundamentals, it has some limitations. The MS unit root test relies on the researcher's discretion to decide whether explosive behaviour exists at the same time in fundamentals.

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[47] Note that such analysis is not applicable to emerging market economy countries as their exchange rate regimes are, in general, fixed or managed regimes (crawling peg) for the period under examination. This condition does not allow for an accurate analysis of bubbles. Low movements of exchange rates may usually be attributed to eventual speculative attacks rather than the presence of bubbles.

Deciding whether or not exchange rates depart in an explosive way from fundamentals is not testable with unit root tests. For this reason, this chapter uses an additional econometric approach to test for bubbles, based on Markov-switching vector autoregressions (MS-VECM) proposed by Krolzig (1996). The MS-VECM was originally applied to business cycle analysis and allows investigation of the existence of long-run relationships with regime switching. In essence, such a procedure contains an error correction mechanism which corrects deviations from long-run relationships, and takes into account dynamic adjustments concerning the transition from one regime to another. The importance of using the MS-VECM approach to investigate the presence of bubbles in the stochastic process is that tests applied to time series with uncontrolled for regime switching may lead to misleading conclusion in favour of the presence of bubbles. Hence, this approach allows the analysis of long-run behaviour of time series by eliminating the effect of regime switching on the stochastic process. Artis and Krolzig (2004), Krolzig and Toro (2001) and Krolzig and Sensier (2000) are recently published studies related to the business cycle which make use of MS-VAR modelling. It is still important to emphasise that the employment of the MS unit root test and the MS-VECM approach is new in the empirical literature on the analysis of bubbles in exchange rates.

The chapter is structured in five distinct sections. Section 4.1 outlines the main points of the theoretical model to be employed. Section 4.2 describes the statistical data, variables used and the research sources. Section 4.3 presents a theoretical overview of the Markov-switching regime unit root test and the Markov-switching vector autoregression-based cointegration test as well as the empirical results. Finally, Section 4.4 presents the conclusions.

4.1 - The Theoretical Model

The model used in this chapter follows a standard monetary model of exchange rate determination, which derives the general solution for bubbles under rational expectations. It follows similar versions used by Evans (1991), Taylor and Peel (1998), Hall *et al.* (1999), Hooker (2000), Psaradakis *et al.* (2001). The model of two countries, domestic and foreign, assumes each country has a money demand equation with the same income elasticity and interest rate semi-elasticity. Thus, it is possible to obtain a single money market equilibrium by combining the money market equilibrium condition in both countries. It leads to the following specification:

$$m_t - p_t = a_1 y_t - a_2 i_t \qquad (4.1.1)$$

where m_t, p_t, y_t and i_t denotes, respectively, the natural logs of the relative money supply, the relative price level, relative real income and relative interest rate between the domestic and foreign countries. The coefficients a_1 and a_2 correspond to the income elasticity and the interest rate semi-elasticity, respectively.

As the variable i_t in (4.1.2) denotes the interest rate differential between the countries and it is assumed that the uncovered interest rate parity (UIRP) condition holds the following expression emerges:

$$i_t = E_t e_{t+1} - e_t \qquad (4.1.2)$$

where e_t denotes the natural log of the nominal exchange rate at time t and E_t is the mathematical expectation condition on information available at time t. Equation (4.1.2) means that an eventual non-zero interest rate differential must be offset by an equivalent expected change in the exchange rate..

The model also assumes that deviations from purchasing power parity (PPP) follow a random walk given by:

$$e_t - p_t = u_t \qquad (4.1.3)$$

where $u_t = u_{t-1} + \varepsilon_t$ and $\varepsilon_t \sim N\left(0, \sigma_\varepsilon^2\right)$. Assuming that m_t and y_t are exogenous and substituting (4.1.2) and (4.1.3) into (4.1.1) results in the following representation:

$$e_t = (1-b)f_t + bE_t e_{t+1} + (1-b)u_t \qquad (4.1.4)$$

$$\text{and } 0 < b \equiv \frac{a_2}{1+a_2} < 1$$

where $f_t = m_t + a_1 y_t$ denotes the market fundamental solution.

From equation (4.1.4) a first order expectational difference equation may be derived by repeatedly substituting $E_t e_{t+1}$ for n future time periods. As a result of this, when $j \to \infty$ and $(1+a_2)^{-1}$ is less than unity, by hypothesis, a non-bubble solution e_t^f emerges as follows:

$$e_t^f = (1-b)\sum_{j=0}^{\infty} b^j E_t f_{t+j} + u_t \qquad (4.1.5)$$

However, the difference equation (4.1.5) can also have an infinite number of solutions if $(1+a_2)^{-1}$ is greater than unity:

$$e_t^f = (1-b)\sum_{j=0}^{\infty} b^j E_t f_{t+j} + B_t \; = \; e_t^f + B_t \qquad (4.1.6)$$

Thus, if $\{B_t\} \neq 0$, and it is assumed to be an implicit process of (4.1.6) and defined by (4.1.7) as:

$$E_t\left(B_{t+j}\right) = B_t\left(\frac{1}{b}\right)^j, \qquad for \quad j = 0,1,2,... \qquad (4.1.7)$$

then the solution to (4.1.5) is not unique and a potential infinite set of solutions derives from (4.1.7).

Therefore, the solution associated with (4.1.5) is the market fundamental solution and (4.1.6) is a whole set of bubble solutions in which B_t is the exchange rate bubble. The extent of the deviation of the exchange rate from the market equilibrium is a rational bubble captured by the term B_t defined by (4.1.7), which drives exchange rate movements away from the market fundamental solution. If such a deviation is perceived by market participants to be significant for speculative purposes, then it will be assigned probabilities and a data generating process (DGP) is formed. The DGP delivers the actual rational aspect to a rational bubble. There will be a particular probability associated with the continuation of the bubble next period against the probability of the bubble bursting. It is worth emphasising that given the bubble's asymmetric probability distribution, then the distribution of the exchange rate innovations will also be asymmetric.

In the real world, this process cannot only be thought of as being characterised by deterministic bubbles. If it is assumed that bubbles do exist, they must be a stochastic or a periodic process, so that there are periods of expansion as well as periods of contraction or collapse. Evans (1991) embraces this idea, and demonstrates the existence of an important class of periodically collapsing bubbles which is a key focus of this chapter. It can be described according to the following representation:

$$B_{t+1} = \begin{cases} (1+r)B_t u_{t+1} & if \quad B_t \leq \alpha \\ \left[\delta + \pi^{-1}(1+r)\theta_{t+1}\left(B_t - (1+r)^{-1}\delta\right)\right]u_{t+1} & if \quad B_t > \alpha \end{cases} \tag{4.1.8}$$

where $(1+r)$ is the discount rate, assumed to be constant, δ and α are real positive parameters such that $0<\delta<(1+r)\alpha$, $\{u_t\}$ is a sequence of non-negative exogenous i.i.d. random variables with $E_t u_{t+1} = 1$, and θ_{t+1} is an exogenous i.i.d Bernoulli process independent of $\{u_t\}$, such that it takes the value 1 with probability π and 0 with the probability $(1-\pi)$ and $0<\pi\leq 1$. Note that δ is the mean value of a bubble and α is a positive parameter denoting the magnitude of the bubble, from which a bubble can take on a new dynamic. The idea is that when $B_t > 0$ it implies that $B_s > 0 \quad \forall s > t$. If $B_t \leq \alpha$ the bubble will be growing at mean rate $(1+r)$. On the other hand, when $B_t > \alpha$ the bubble acquires a new expansion dynamic at faster mean rate of $(1+r)\pi^{-1}$ until the bubble collapses with probability $(1-\pi)$. Once the bubble has collapsed it restarts and expands from the mean value of δ.

4.2 - Data

The data used in this chapter are collected from the International Financial Statistic-IFS provided by the International Monetary Fund with end-of-quarter periodicity. They consist of four industrialised market economy countries: Canada, France, Germany and the United Kingdom. The period of analysis extends from 1973:1 to 1998:4, since from 1999:1 onwards a single currency system (the Euro) was introduced.

This study conducts the analysis by using the nominal exchange rate (ER) against the US dollar country by country as the endogenous variable. The regressor denoting money supply consists of the narrow money supply differential $M1$ ($m1 - m1^* = M1$). The regressor is the real income differential ($gdp - gdp^* = Y$), and real income is represented by GDP in volume.

Finally, the variable differentials are calculated by assuming the United States as the *numeraire* country denoted by an asterisk and in log forms.

4.3 – Econometric Methodology

4.3.1 - The Markov-Switching Unit Root Test

A number of papers has demonstrated that tests of unit roots and cointegration for economic series may fail in the presence of periodically collapsing bubbles (see Evans 1991). Using Monte Carlo simulations Evans (1991) showed that economic time series with a bubble process may often appear to be stationary in terms of traditional unit root and cointegration tests, even though bubbles are explosive by construction. Taylor and Peel (1998) argue that since bubbles must collapse periodically, standard tests for unit roots and cointegration can generate the misleading conclusion of monotonic mean reversion.

Given these weakness of the familiar augmented Dickey-Fuller (ADF) unit root test in the presence of periodically collapsing bubbles in economic time series, this chapter makes use of a more robust test which was explored by Hall *et al.* (1999)[48] and extends it to a multivariate context. The popular approach to constructing tests of hypotheses to detect the presence of a unit root in time series $\{y_t\}_{t=1}^{n}$ is based on autoregressive regression models like equation (4.3.1.1):

$$\Delta y_t = \mu + \phi y_{t-1} + \sum_{j=1}^{k} \psi_j \Delta y_{t-j} + v_t \qquad (4.3.1.1)$$

48 Note that a similar procedure based on Markov regime-switching regression models was also employed by Van Norden (1996, 1998).

where $\Delta y = y_t - y_{t-1}$, $\{v_t\}$ is a zero-mean white noise process, and κ is a suitably chosen integer (see Dickey and Fuller, 1981). The standard analysis is conducted on the coefficient of y_{t-1} (ϕ) in (4.3.1.1), which is examined for its statistical significance based on the t-statistic, but with non-standard critical values. Statistically if $\phi = 0$ the series is supposed to be I(1). Nonetheless, the main difficulty of this class of test is that it is only able to detect the presence of a bubble in its expansion phase.

A generalisation to the standard ADF unit root test is to allow for the possibility of the dynamic behaviour of $\{y_t\}$ assuming different characteristics for different periods of the sample. Making use of the class of dynamic Markov-switching models explored in Hamilton (1988, 1989), Hall *et al.* (1999)'s approach consists of modifying the standard ADF unit root test allowing the ADF regression parameters in equation (4.3.1.1) to switch values over different regimes/states. This procedure allows for a dynamic structure consistent with periodically collapsing bubbles. The motivation for this new methodology is the possible existence of two different regimes driving the economic series congruent with the expanding and collapsing phases of the bubble. The generalisation consists of assuming that the parameters governing equation (4.3.1.1) are time-varying, that is, they change with an unobserved indicator $s_t \in \{0,1\}$, so that the generalised equation can be denoted according to the following representation:

$$\Delta y_t = \mu_0(1-s_t) + \mu_1 s_t + \left[\phi_0(1-s_t)+\phi_1 s_t\right] y_{t-1} + \sum_{j=1}^{k}\left[\psi_{1j}(1-s_t)+\psi_{2j} s_t\right]\Delta y_{t-j} + e_t \qquad (4.3.1.2)$$

where $\{e_t\}$ is a sequence of independent and identically distributed (i.i.d) random variables with zero mean and unit variance.

According to Goldfeld and Quandt (1973) and Hamilton (1988, 1989), date t is associated with a regime selected naturally and with a probability which depends upon what regime preceded the process at date t-1. The idea is that given a random sequence $\{s_t\}$, characterised as a homogeneous Markov chain on the state space $\{0,1\}$, the transition probability associated with each different state space is:

$$\begin{aligned}
&\Pr(s_t = 1 \mid s_{t-1} = 1) = p \\
&\Pr(s_t = 0 \mid s_{t-1} = 1) = 1 - p \\
&\Pr(s_t = 0 \mid s_{t-1} = 0) = q \\
&\Pr(s_t = 1 \mid s_{t-1} = 0) = 1 - q
\end{aligned} \qquad (4.3.1.3)$$

where an additional requirement is that the innovations $\{e_t\}$ in equation (4.3.1.2) must be independent of the state variables $\forall_t$. By allowing the model's parameters to be functions of the stochastically chosen regimes which control the process at date t, the equations (4.3.1.2) and (4.3.1.3) constitutes a generalisation of the linear ADF model in (4.3.1.1). It is important to highlight that these regimes or states changes allow a variety of outcomes to take place.

Like the traditional ADF unit root test, the test criteria for the null hypothesis of a unit root in either regimes requires that the parameters $\phi_0 = 0$ and/or $\phi_1 = 0$. Hence, by analogy to the standard ADF statistic, the Markov-switching unit root test may be based on the asymptotic t-ratios associated with the maximum likelihood estimates of ϕ_0 and ϕ_1. Nevertheless, detecting the presence of an explosive rational bubble within the Markov-switching ADF framework requires $\phi_0 > 0$ or $\phi_1 > 0$, indicating that one of the regimes driving the stochastic process has an explosive autoregressive root[49].

The MS-switching regime unit root test is applied for the exchange rate and its fundamentals in four industrialised market economy countries: Canada, France, Germany and the United Kingdom. As the fundamentals model of the exchange rate corresponds to the traditional monetary approach, the tests are applied to the exchange rate, money supply differential and output differential. Note that the uncovered interest rate parity (UIRP) condition holds. It is also assumed that the time series contain two different regimes: one of them represents the long-run fundamentals solution and the other one represents the non-fundamentals solution. The non-fundamentals solution is supposed to characterise a bubble process. It is important to emphasise that evidence of explosive behaviour found in the exchange rate series combined with similar behaviour in its fundamentals series is suggestive that exchange rate volatility is a consequence of market fundamentals volatility. On the other hand, if exchange rate behaviour is not followed by similar behaviour in its fundamentals then a rational bubble process may be driving the exchange rate away from the fundamentals solution.

In practice, the MS-unit root test procedure composed of two different regimes consists of testing the null hypothesis of non-stationarity $\phi_0 = 0$ and $\phi_1 = 0$ against the alternative $\phi_0 < 0(\phi_1 > 0)$ and $\phi_0 > 0(\phi_1 < 0)$ in equation (4.3.1.2). Note that a negative estimate ϕ_s and a positive estimate ϕ_s statistically significant in the regime

[49] It is worth emphasising, as stated by Hall et al. (1999), that the Markov-switching unit root test is a methodology designed to detect successfully the presence of all types of periodically collapsing bubbles. However, if the probability of the bubble collapsing is significantly high, or if the effect of a bubble driving the volatility of asset prices is not relevant, then the power of any bubble test becomes very low.

$s_t = \{0,1\}$ imply stationarity and explosive behaviour, respectively. An explosive estimate of ϕ_s for the exchange rate in the regime s_t, not followed by similar estimates in its fundamentals, indicate the possibility of bubbles governing the stochastic process. Table 4.1 displays the results of estimating (4.3.1.2) for Canada, France, Germany and the United Kingdom. The lag order of the ADF regression is based on the Akaike information criterion[50].

[50] Note that all computations are generated by a regime-dependent coefficients and heteroskedasticity MSIAH(s_t)-ARX(p_t) model with two different regimes and p_t lag order for variables in first difference.

Table 4.1 – Maximum Likelihood Estimates for MS-ADF Regression[1]

Country	Variable[51]	ϕ_s	
		$s_t = 0$	$s_t = 1$
Canada[2]	Exchange Rate [4]	0.1895 (23.884)**	-0.0261 (-1.347)
	Money Supply [4]	0.1621 (15.315)**	-0.0918 (-2.995)**
	Output[4]	-0.1009 (-12.519)**	0.0142 (0.518)
France[2]	Exchange Rate [3]	-0.1169 (-3.184)**	-0.0227(-0.723)
	Money Supply [1]	-0.3302 (-19.568)**	-0.0371 (-1.289)
	Output [2]	-0.0914 (-3.755)**	0.0513 (4.238)**
Germany[2]	Exchange Rate [5]	-0.1101(-3.378)**	-0.0296 (-0.825)
	Money Supply [1]	0.0119 (0.3704)	-0.0898(-2.067)
	Output [4]	-0.053 (-2.585)*	0.015 (0.075)
UK[2]	Exchange Rate [3]	-0.1889 (-4.519)**	-0.1235 (-2.539)*
	Money Supply [4]	-0.1595 (-4.217)**	-0.0052 (-0.139)
	Output [1]	-0.0773(-3.795)**	0.1004 (3.77)**

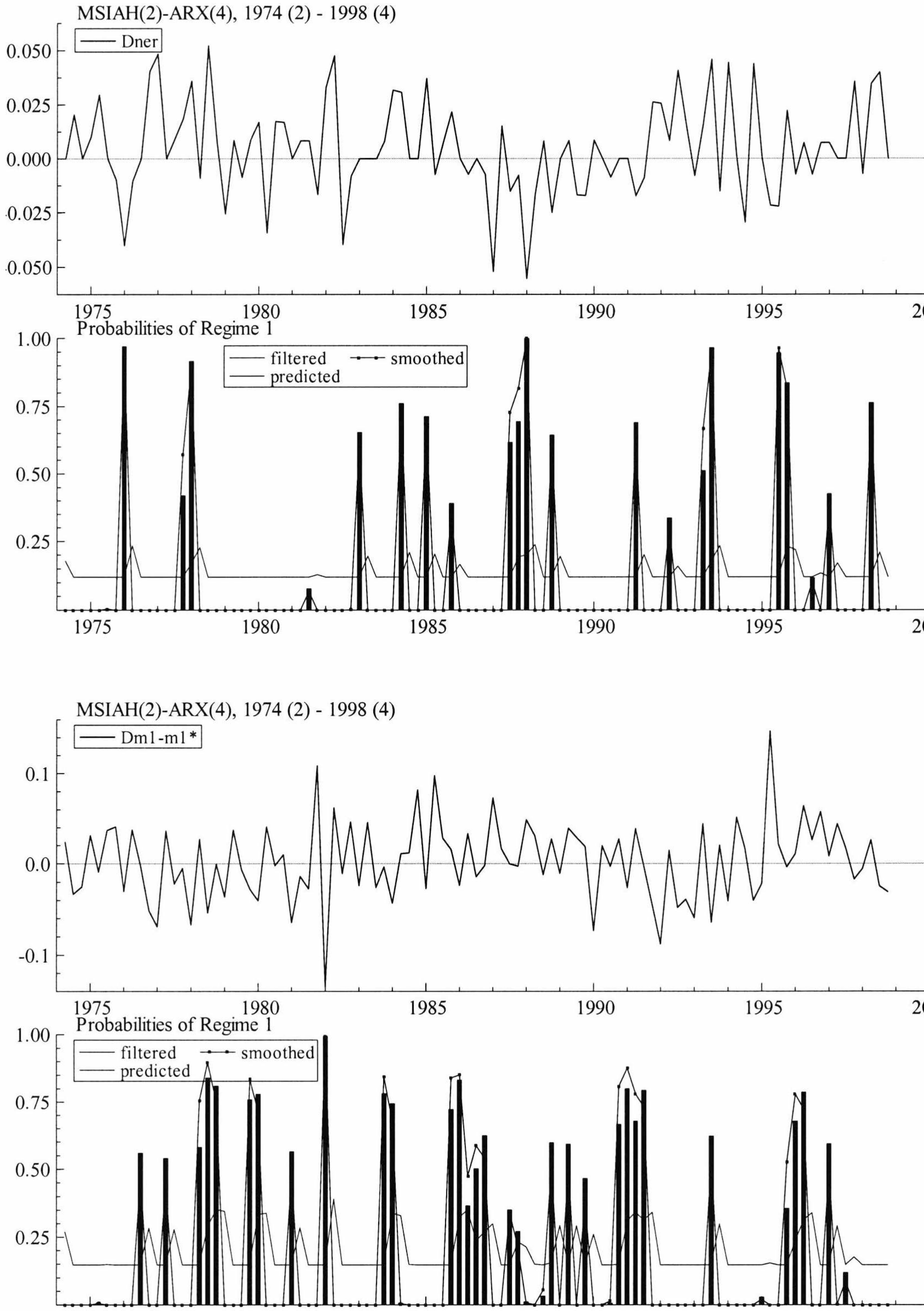
[1] Figures in square bracket are the lag order in the ADF regression and those in parentheses are *t*-values.
[2] Critical Values: * = (5% level) , ** = (1% level).

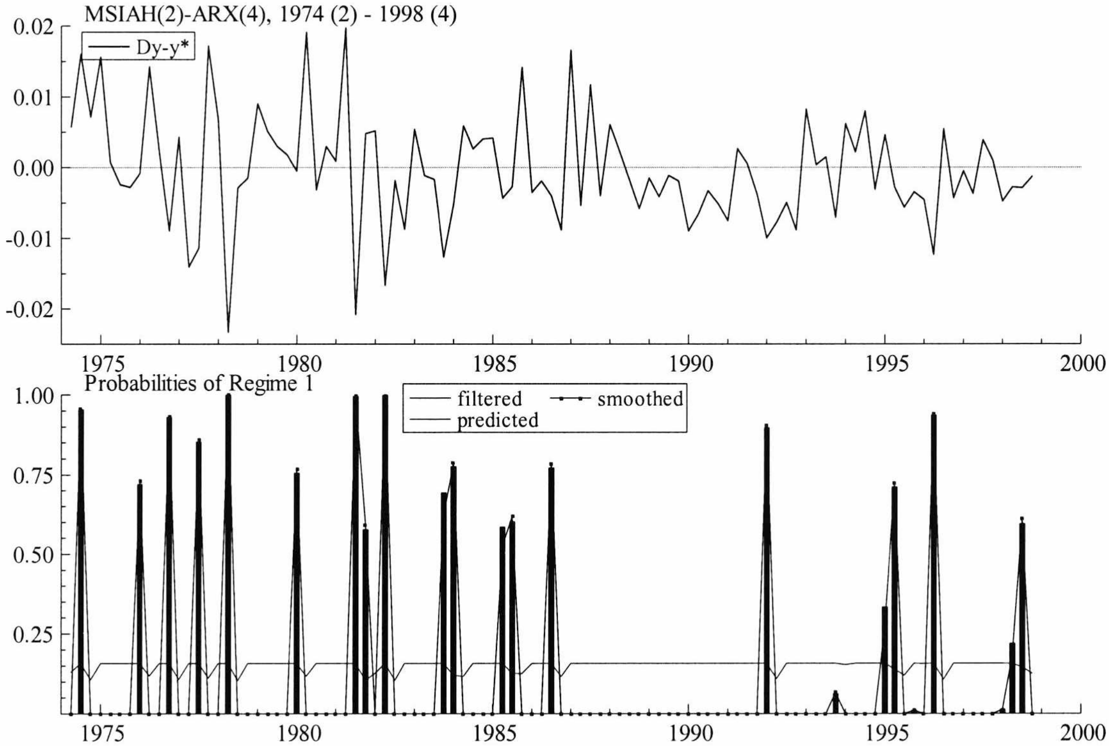
The results in Table 4.1 for Canada show a statistically significant positive estimate of ϕ_0 ($s_t = 0$) and a non-significant negative estimate of ϕ_1 ($s_t = 1$) for the exchange rate. Note, nevertheless, that the computation for the money supply also demonstrates a significantly explosive result of ϕ_0 in regime $s_t = 0$.

[51] Note that all variables are expressed in logarithmic form.

Thus, the positive estimate in regime $s_t = 0$ for both variables rejects the hypothesis of bubbles driving the exchange rate. As the exchange rate and the money supply demonstrated both explosive behaviours in $s_t = 0$, the hypothesis of the presence of bubbles is rejected in favour of a fundamental solution. This conclusion for Canada is partially reinforced by analysing the probabilities considering the sample as a whole in regimes $s_t = 0$ and $s_t = 1$ (see Figure 4.1-Canada). The probability of the exchange rate movements remaining in regime $s_t = 0$ associated to the probability of the money supply remaining in the regime $s_t = 0$ rejects the presence of bubbles. Nonetheless, in some quarters before and after 1980 and 1990 the exchange rate demonstrates an explosive behaviour as in regime $s_t = 0$ associated to non-explosive fundamentals evidenced in regime $s_t = 1$. The graphical analysis for these periods may reveal some evidence of bubbles and requires further investigations.

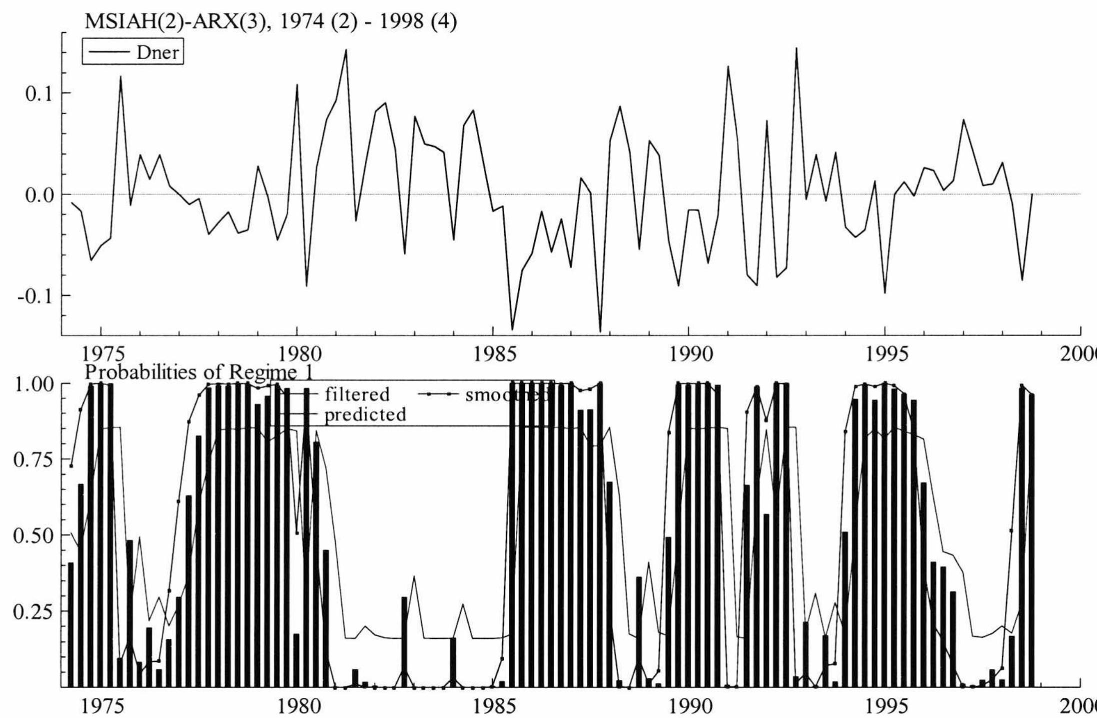
Figure 4.1-Canada

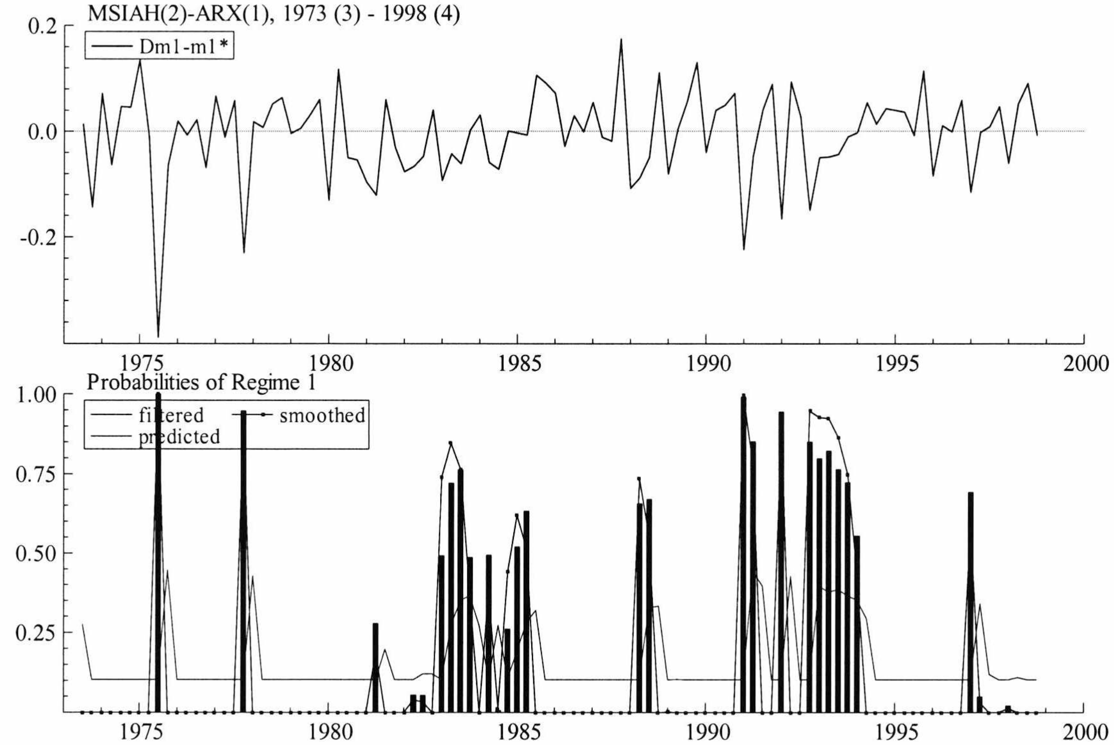


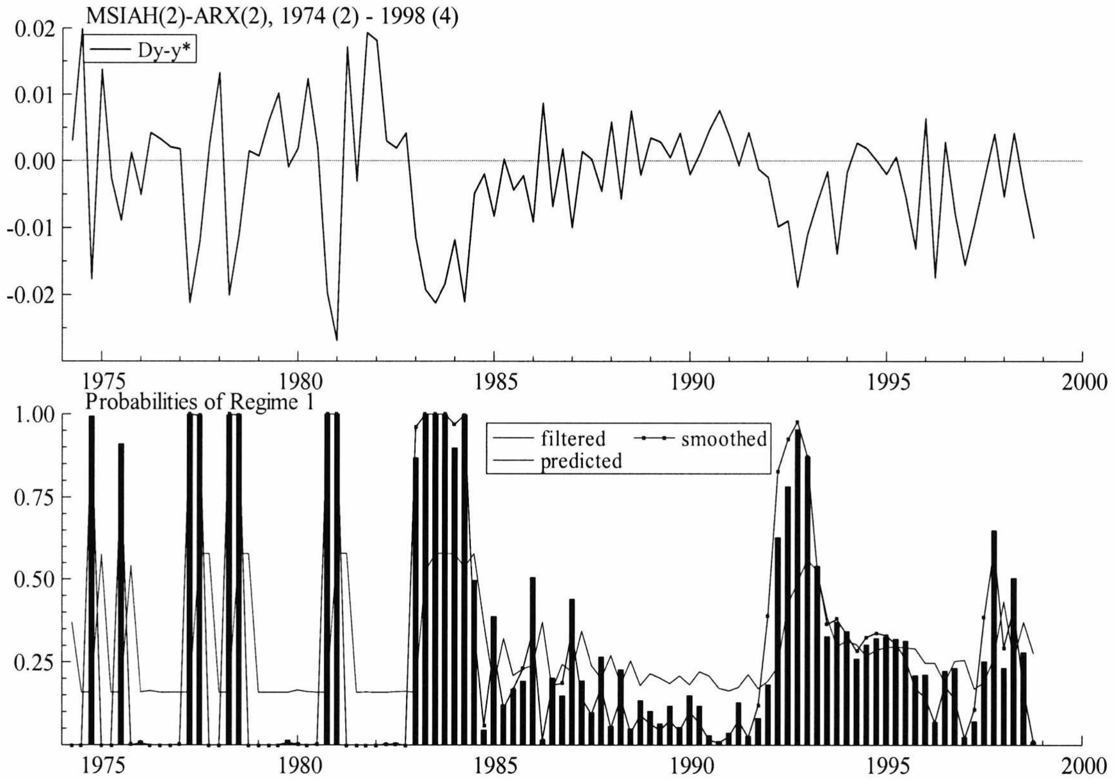


The analysis for France is clear cut as the negative estimates of ϕ_s found in both regimes ($s_t = 0$ and $s_t = 1$) for the exchange rate do not support the presence of bubbles (see Table 4.1). The significant negative estimate of ϕ_0 for the exchange rate and for the money supply shows stationarity for both variables in regime $s_t = 0$. The regime $s_t = 1$ produced a non-significant negative estimate of ϕ_1 for the exchange rate and for the money supply. Hence, these results may reveal that the exchange rate behaviour is due to the money supply behaviour and so the hypothesis of the presence of bubbles is again rejected. Figure 4.2 – France shows that the stationarity observed in the exchange rate for the 1981-1985 period and for the 1991-1993 period described in regime $s_t = 0$, in particular, are associated to a mix of non-stationary and stationary behaviours generated by the money supply in regime $s_t = 0$ and $s_t = 1$, respectively. Once again, the graphical analysis reinforces the hypothesis of no bubble driving the exchange rate. Figure 4.2-France displays the graphs of regime probabilities.

Figure 4.2-France

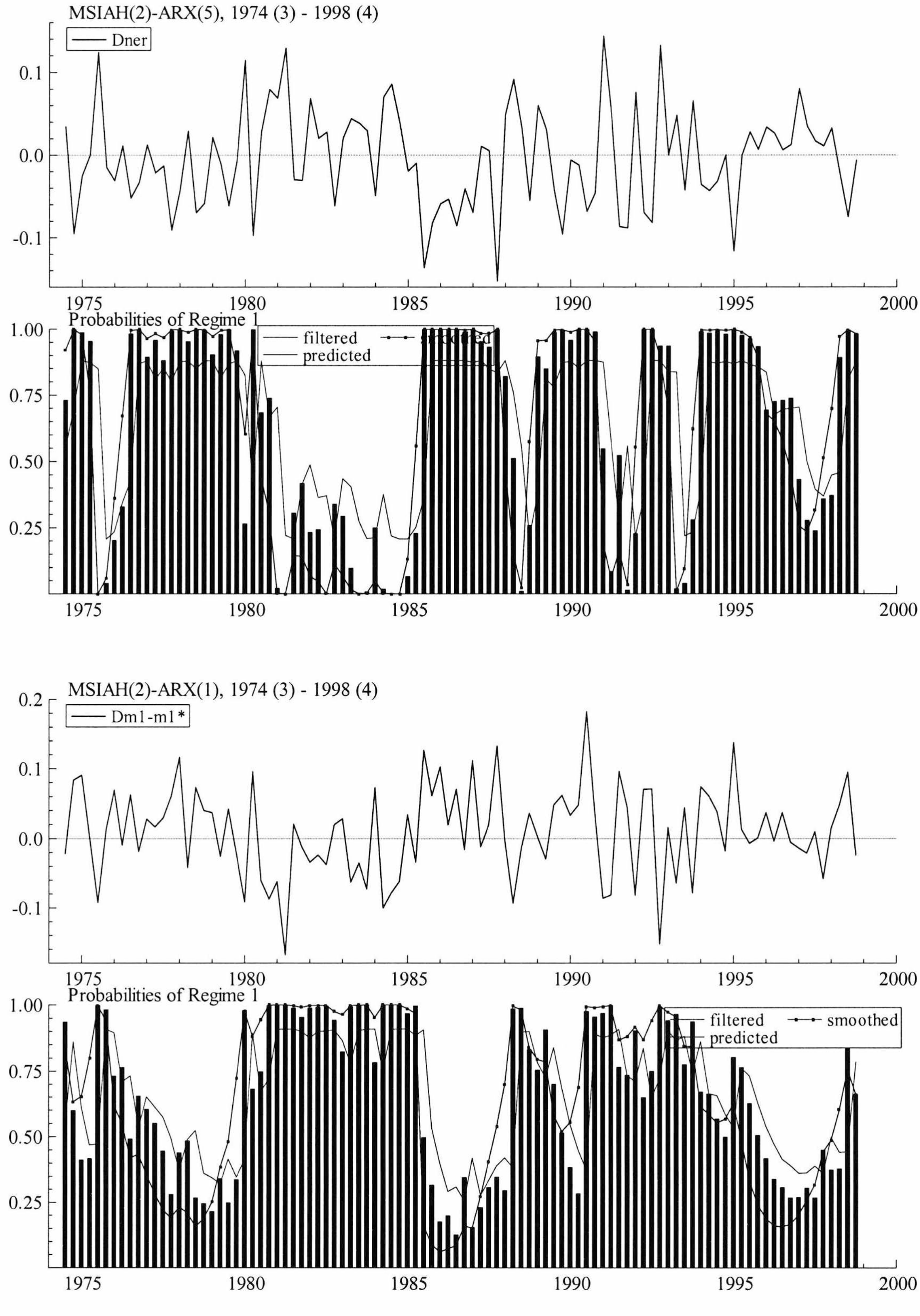


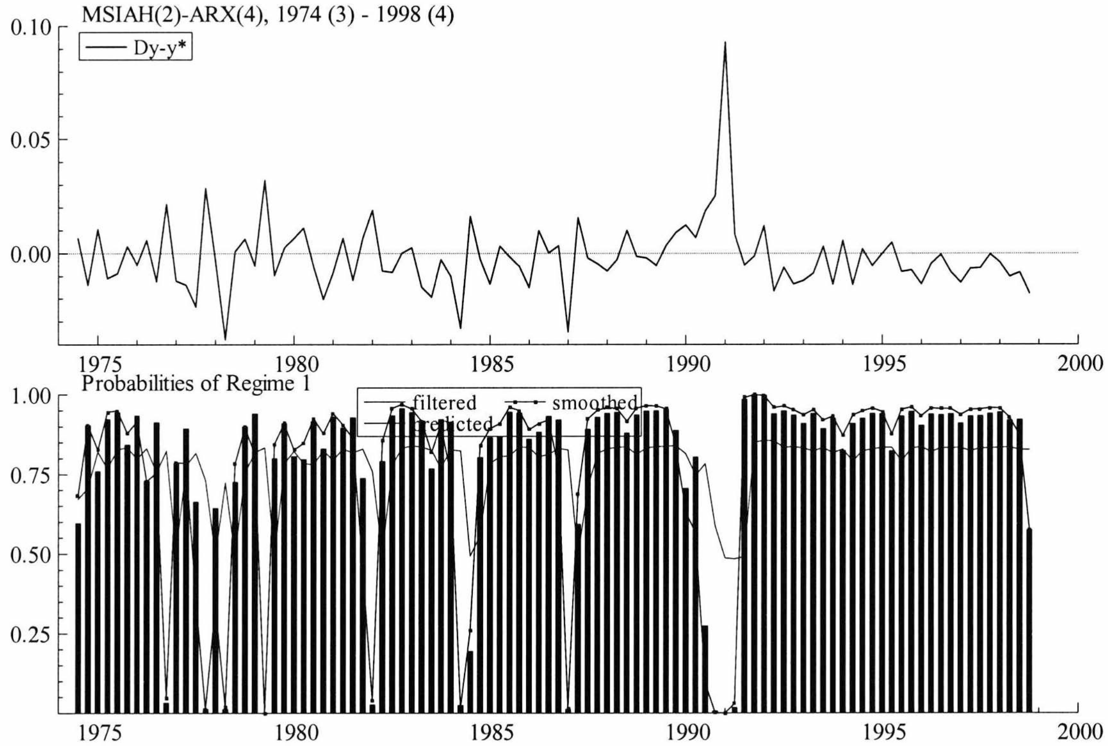

MSIAH(2)-ARX(1), 1973 (3) - 1998 (4)
Dm1-m1*
0.2
0.0
-0.2
1975
1980
1985
1990
1995
2000
Probabilities of Regime 1
filtered
smoothed
predicted
1.00
0.75
0.50
0.25


MSIAH(2)-ARX(2), 1974 (2) - 1998 (4)
Dy-y*
0.02
0.01
0.00
-0.01
-0.02
1975
1980
1985
1990
1995
2000
Probabilities of Regime 1
filtered
smoothed
predicted
1.00
0.75
0.50
0.25

For Germany a positive estimate of ϕ_s is not computed in both regimes for the exchange rate (see Table 4.1). The significant negative estimate of ϕ_s found in regime $s_t = 0$ and non-significant estimate found in regime $s_t = 1$ leads to the conclusion that the exchange rate series is a stationary and non-stationary process, respectively, in different regimes. The estimate of ϕ_s for the money supply is found positive in regime $s_t = 0$, though statistically non-significant, and non-stationary in regime $s_t = 1$. As such, the hypothesis of bubbles in the German exchange rate is rejected for the sample period. This conclusion is supported by analysing graphically the regime probabilities in which the non-stationarity for the exchange rate and for the money supply observed in regime $s_t = 1$ prevails for the sample period as whole (see Figure 4.3-Germany).

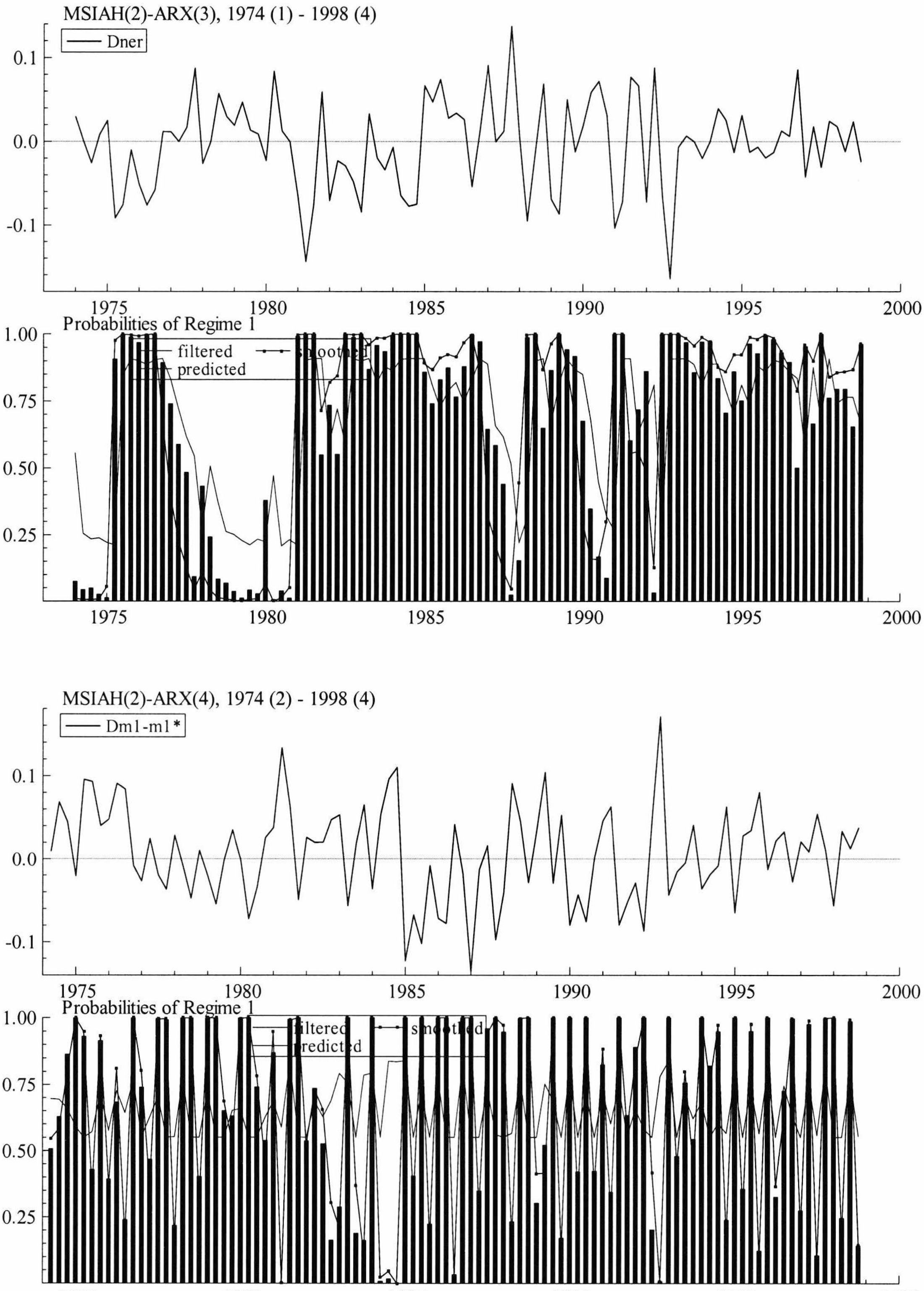
Figure 4.3-Germany

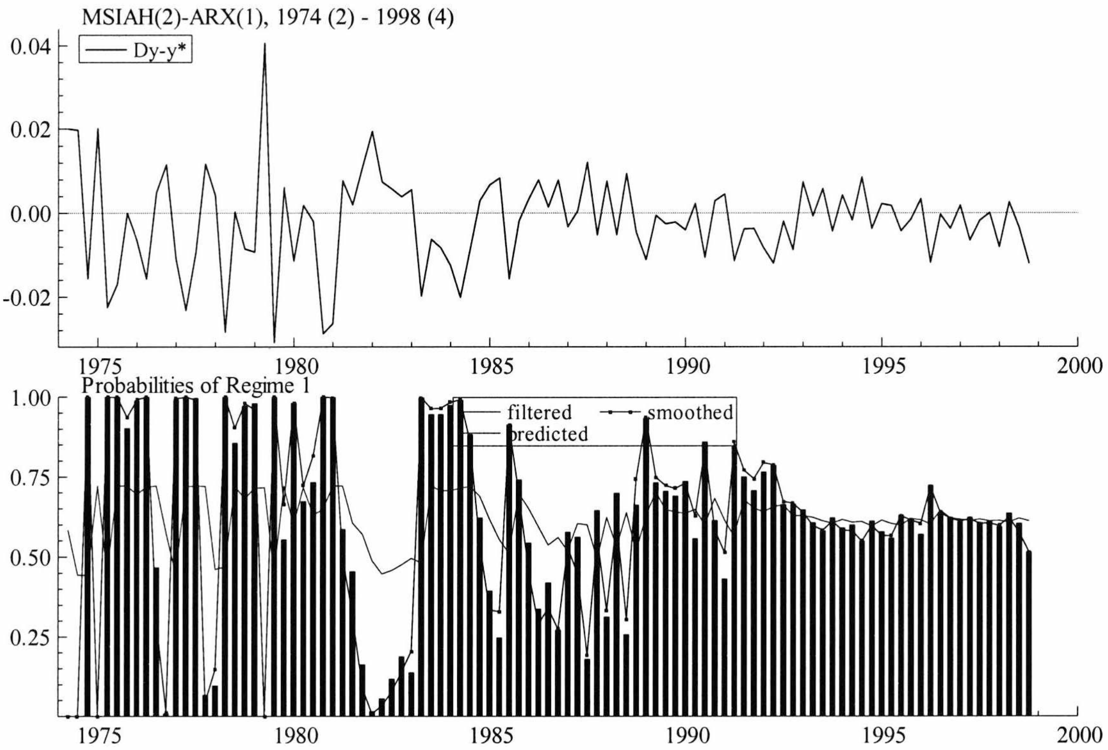




Finally, the United Kingdom case demonstrates a similar conclusion as for Germany about the presence of bubbles. The significant negative estimates of ϕ_s in regime $s_t = 0$ confirm that the exchange rate is a stationary process reinforced by a similar result for the money supply (see Table 4.1). The regime $s_t = 1$ demonstrates that the exchange rate is non-stationary associated to a similar behaviour for the money supply. The results for both regimes do not allow for the conclusion of bubbles driving the exchange rate. Figure 4.4-UK that displays the probabilities of regimes confirm this finding as the graphical behaviour for the exchange rate and the money supply follows the non-stationary regime $s_t = 1$.

Figure 4.4UK


MSIAH(2)-ARX(3), 1974 (1) - 1998 (4)
Dner
0.1
0.0
-0.1
1975
1980
1985
1990
1995
2000
Probabilities of Regime 1
1.00
0.75
0.50
0.25
filtered
smoothed
predicted
MSIAH(2)-ARX(4), 1974 (2) - 1998 (4)
Dm1-m1*



Although it is not possible to identify robustly different regimes (non-bubble and bubble regimes) based on only regime probabilities, it is possible to state the prevailing regimes for each variable. Table 4.2 displays the regime probability for each variable based on the number of observations by regimes.

Table 4.2 – Regime Probabilities (%)*

Country	Variable[52]	$s_t = 0$	$s_t = 1$
Canada	Exchange Rate [4]	13.6	86.4
	Money Supply [4]	19.5	80.5
	Output[4]	15.0	85.0
France	Exchange Rate [3]	51.7	48.3
	Money Supply [1]	83.8	16.2
	Output [2]	72.2	27.8
Germany	Exchange Rate [5]	63.4	36.6
	Money Supply [1]	64.6	35.4
	Output [4]	81.9	18.1
UK	Exchange Rate [3]	69.1	30.9
	Money Supply [4]	65.1	34.9
	Output [1]	61.4	38.6

* Figures in square bracket are the lag order in the ADF regression as in Table 4.1.

By comparing results in Table 4.1 and Table 4.2 for Canada, for instance, it may be interpreted that the regime probability of the explosive behaviour running the exchange rate and the money supply in $s_t = 0$ is much less than the probability for the non-explosive regime $s_t = 1$. It means that a fundamental solution prevails for the explosive case[53]. France presented the regime

[52] Note that all variables are expressed in logarithmic form.
[53] An exception must be made for some quarters before and after 1980 and 1990 as already aforementioned.

probability of a stationary behaviour in regime $s_t = 0$ for the exchange rate and the money supply slightly greater than the regime probability for a non-stationary regime $s_t = 1$. Once again, the fundamental solution for France prevailed. Germany and UK did not exhibit an explosive behaviour for either the exchange rate or the money supply and the prevailing regime probability is stationary. The unexpected regime probabilities found for Germany and UK supporting the stationarity for the exchange rate may require further investigations.

Finally tests for linearity based on MS-likelihood ratios rejected the null hypothesis of one regime and revealed non-linear behaviours for the exchange rate. Table 4.3 displays linearity statistics:

4.3 - Likelihood Ratio (LR) Linearity Tests

Country	**LR**	χ^2
Canada	30.804	(7)[0.0001]**
France	13.5779	(6)[0.0347]*
Germany	22.459	(8)[0.0041]**
UK	12.0129	(6)[0.0417]*

**** = 1% and * = 5% level of significance.**
Figures in parentheses denote degrees of freedom and in squared brackets are *p*-values.

To summarise, the analysis to detect the presence of bubbles based on the MS-unit root procedure does not reveal clear evidence of bubbles driving the stochastic process of the exchange rate. The result for Canada was the only one that produced significant positive estimate of ϕ_s in the exchange rate followed by similar results in at least one exchange rate fundamental. The graphical analysis for Canada also presented an explosive regime in the exchange rate associated to a non-explosive regime in fundamentals which may suggest evidence of bubbles. However, this conclusion is not robust and requires a further investigation. Moreover, as tests detected non-linear behaviours there may be periods in graphs demonstrating some evidence of regime mismatches between the exchange rate and its fundamentals. These periods may require a further investigation. Next section presents an additional approach for bubble tests based on MS-VECM.

4.3.2 - The Markov-Switching Vector Autoregression of Nonstationary TS

The Hamilton (1988, 1989) approach models regime switching behaviour of univariate time series. In particular, the underlying idea of this class of regime-switching models consists of examining the effect of an unobservable regime variable s_t, denoting different states of the world, on the parameters of a stochastic process. The increasing interest in analysing multivariate systems with regime switching encouraged the development of a new technique based on vector autoregression (VAR) modelling. This new approach was proposed by Krolzig (1996) who developed the idea of cointegrated Markov-Switching vector autoregression (MS-VAR). In particular, it is designed to investigate the statistical properties of multivariate time series subject to regimes shifts. Note that although the parameters are time-varying, they are constant conditional on s_t. Essentially, Krolzig extends the univariate case to the multivariate case by generating the finite order vector autoregressive (VAR) model with regime switching which is expressed generically by the following equation:

$$x_t = v(s_t) + A_1(s_t)x_{t-1} + \ldots + A_p(s_t)x_{t-p} + \mu_t \qquad (4.3.2.1)$$

where the sample values $x_0, \ldots, x_{1-p}$ are fixed and constitute elements of a vector matrix. The dependence of the parameters on the realised regime s_t is denoted by the parameter shift functions $v(s_t)$, $A_1(s_t), \ldots, A_p(s_t)$. For instance, the state-space representation for the intercept is expressed as:

$$v(s_t) = \begin{cases} v_1 & if\ s_1 = 0 \\ \vdots & \\ v_M & if\ s_t = M-1 \end{cases} \qquad (4.3.2.2)$$

The parameters in equation (4.3.2.1) are subject to a prevailing regime which is stochastic and unobservable. Thus, a complete description of the data generating process requires formulation of a rule, a regime generating process, and then the evolution of regimes can be inferred from the data. The regime generating process in Markov-switching models constitutes a homogeneous Markov chain based on a finite number of states $s_t = 1, \ldots, M$ and defined by transition probabilities:

$$p_{ij} = P_r(s_{t+1} = j \mid s_t = i), \quad \sum_{j=1}^{M} p_{ij} = 1 \quad \forall_{i,j} \in \{1, \ldots, M\} \qquad (4.3.2.3)$$

The idea of a MS-VAR process is based on the existence of a finite-order vector moving average (VARMA) representation[54]. Such a VARMA structure may be approximated by a finite-order linear VAR model, and estimated by Johansen's maximum likelihood procedure (see Johansen, 1995).

The stationarity of a time series requires that the mean and variance are constant (unconditional). Stylised facts show that many economic time series exhibit non-stationary behaviour but it is possible for existing linear relations between non-differenced variables to lead to stationarity. If there is at least one linear combination of variables ($c'y_t$) which is stationary, then the I(1) process is said to be cointegrated. The cointegration relationships are referred to as the long-run equilibrium of the system. Like the standard vector equilibrium correction model (VECM) proposed by Davidson *et al.* (1978) and Engle and Granger (1987), a Markov-switching vector equilibrium correction model (MS-VECM) can also be implemented. The basic idea is that the error correction mechanism contained within a MS-VECM also allows for the errors arising from regime shifts to be corrected towards the stationary distribution of the regimes (see Psaradakis *et al.*, 2004). A generalisation of this model is as follows:

$$\Delta x_t = v(s_t) + \sum_{i=1}^{p-1} \Gamma_i(s_t)\,\Delta x_{t-i} + \Pi(s_t)\,x_{t-p} + u_t \qquad (4.3.2.4)$$

[54] The intercept term is composed of an unconditional mean $\bar{v}$ plus a moving average $\mathrm{MA}(\infty)$ representation.

where $\Gamma_i(s_t) = -\left(\mathrm{I}_k - \sum_{j=1}^{i} \mathrm{A}_j\right)$ is the coefficient matrix on the differenced variables subject to the regime s_t, and $\Pi(s_t) = \alpha\beta' = \mathrm{I}_k - \sum_{j=1}^{p} \mathrm{A}_j$ is the coefficient matrix subject to the regime s_t, which is composed of the adjustment velocity matrix α and the cointegration matrix β'. The matrix $\Pi(s_t)$ corresponds to the error correction mechanism[55].

The rank r of the matrix $\Pi(\alpha\beta')$ defines general conditions for cointegration. If the rank of the matrix Π is $r<k$, then there is a linear combination of variables that is stationary, where k is the number of variables in the model. Furthermore, the regimes are generated by a stationary Markov chain and characterised by attractor elements of the system defined by contemporaneous shifts in the drift $\mu(s_t)$ and in the long-run equilibrium $\delta(s_t)$. The attractors imply an immediate one-time-jump of the process after a change in regime. Such an important stationarity property allows for general effects of regime shifts according to the following representation:

$$\Gamma(L)(\Delta y_t - \mu(s_t)) = \alpha(s_t)\beta'(y_{t-p}) + \mu_t^{56} \qquad (4.3.2.5)$$

In fact, this concept of cointegration is closely related to the concept of co-breaking for multiple time series subject to regime

[55] Note that this correction mechanism is closely related to the concept of a multiple dynamic equilibrium in economics, defined by the equilibrium value of the cointegration vector and the drift.

[56] Equation 4.3.2.5 implies the ECM is not subject to regime shift. The underlying intuition is that the fundamental solution is unique for the exchange rate, therefore the ECM does not admit more than one regime. The ECM is unchangeable.

switching introduced by Hendry (1996). The idea behind the concept of co-breaking consists of removing the effects of regime switching by taking linear combinations of variables. If at least one linear combination remains stationary the co-breaking still prevails even if a regime shift changes the drift of the system. The stationarity of the stochastic process generated by linear combinations implies that the effects of regime switching are dropped asymptotically. The co-breaking condition for a MS-VAR process is given by:

$$\Phi\left[v_1 - v_M, \ldots, v_{M-1} - v_M\right] = \Phi\mathrm{M} = 0 \qquad (4.3.2.6)$$

where M are different regimes and $\Phi = \beta'$ is a $(n \ x \ k)$ matrix composed of n contemporaneous co-breaking combinations.

Similar to the unit root test approach for regime switching, the investigation of the presence of a bubble by a MS(s_t)-VECM(p) requires the analysis of the sign on the adjustment coefficient $\alpha(s_t)$ in equation (4.3.2.5). Once again, assuming that the stochastic process is characterised by two regimes $s_t[0,1]$, one corresponds to the fundamental solution and the other is the bubble solution. The hypothesis of a bubble driving the stochastic process is accepted if $\alpha_0 > 0$ or $\alpha_1 > 0$ are found.

To analyse the presence of bubbles in the exchange rates, a MS-VECM was applied to each country individually according to equation (4.3.2.5). The econometric approach consists, firstly, of estimating a linear VAR with finite order[57]. Next, based on the estimated cointegration matrix, the EM (Expectation-Maximisation)

[57] This procedure aims at estimating the parameters of a linear VAR which comprises the ECM. Once again, a linear VAR is required as there is only one fundamental solution for the exchange rate. The other ones are assumed to be bubble solutions.

algorithm is used to estimate the remaining coefficients of a MS-VECM[58]. In practice, the analysis of a bubble driving the stochastic process consists of examining the significance of the coefficient on $VECM_{t-1}$ in equation 4.3.2.5 which is subject to regime shifts. If it is found to be negative and statistically significant then the process converges towards the fundamental equilibrium, but if it is found to be positive and statistically significant, then it may be interpreted as an explosive process indicating the presence of bubbles[59].

For the case of Canada, a MS(2)-VECM(2) is estimated and the lag order is based on the minimum Akaike information criterion (AIC). The log-likelihood statistic 261.382(240.281), AIC -4.795(-4.632), likelihood ratio test 42.199 all reject the linearity hypothesis of the stochastic process. Even though the Canadian data have produced a positive coefficient on $VECM_{t-1}$ in the regime $s_t = 1$ (see Table 4.3) it is not statistically significant. This may suggest that there is a non-stationary component in the joint exchange rate-fundamentals process, but its statistical insignificance does enable a bubble conclusion. Regime 2 is characterised as non-stationary since the negative coefficient on the ECM is statistically not significant. Also, based on the duration estimates it is possible to observe that the time path for the Canadian exchange rate is remarkably turbulent as regime 1 (duration=11.7) prevails over regime 2 (duration=7.65). The transition probability reinforces this conclusion since the transition probability of leaving the turbulent regime 1 to non-stationary regime 2 is just 8%. Table 4.4 and Figure 4.5-Canada display the main estimated coefficients and the regime probabilities, respectively.

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[58] Note that the VAR estimation procedure has only been applied to the ECM. The other coefficients are estimated in accordance with the Engle and Granger (1987) procedure.

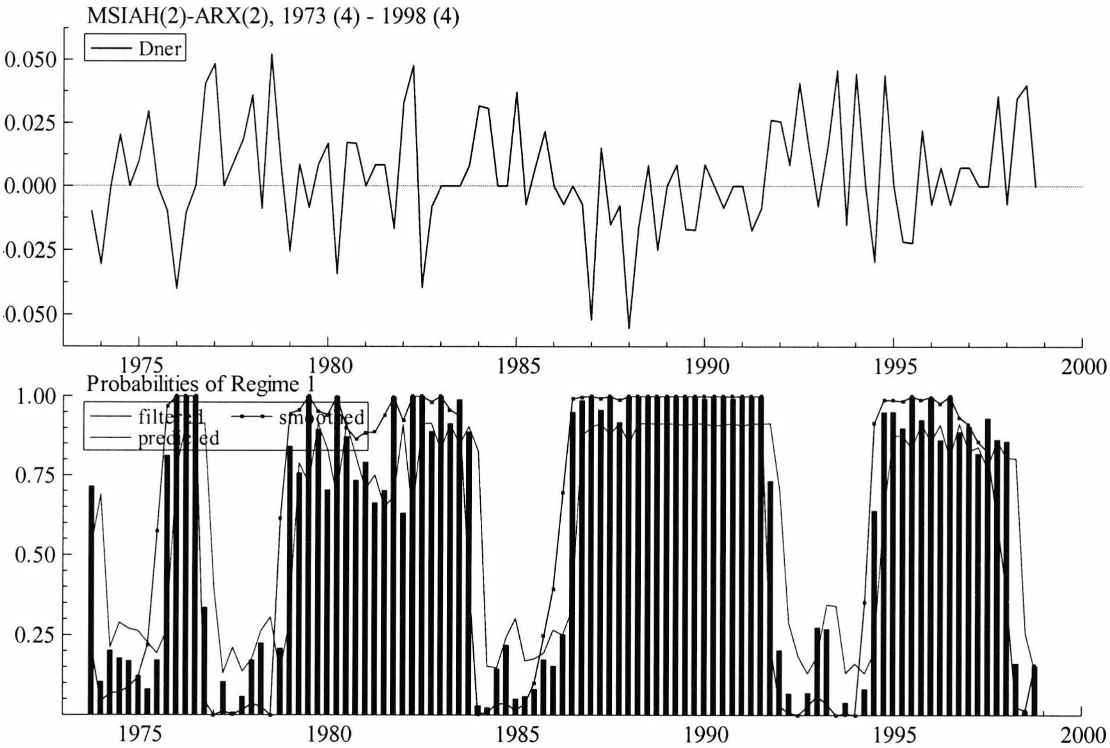
[59] For simplicity, the critical values are taken from the Dickey-Fuller distribution as the distribution for this approach requires a complex simulation process not available for this thesis.

Table 4.4 – Main ML Estimation Results (Canada)[1]

MSIAH(2)-VECM(2)	Regime $s_t = 0$*	Regime $s_t = 1$*
Intercept	-0.0306(-1.1192)	-0.0355(5.662)
VECM$_{t-1}$ $\alpha(s_t)$	0.0104(0.9592)	-0.0181(-1.6215)
Standard Error	0.01592	0.01358
Duration	11.7 (61 obs.)	7.65(39 obs.)
Log-likelihood	261.976 (243.532)	
LR linearity test	36.887, $\chi^2(9) = [0.0000]^{**}$	
Transition Matrix	Regime $s_t = 0$	Regime $s_t = 1$
Regime $s_t = 1$	0.92	0.08
Regime $s_t = 2$	0.13	0.87

* Figures in parentheses denote *t*-statistics.

Figure 4.5-Canada



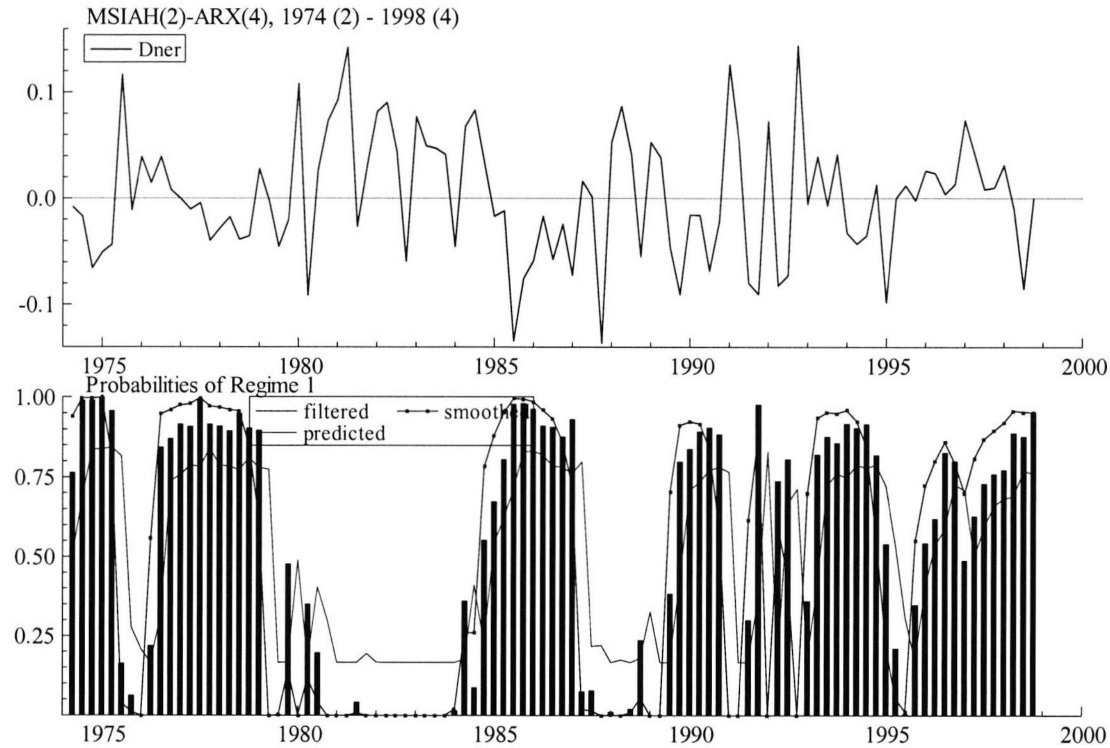
For France a MS(2)-VECM(4) is estimated in accordance with the minimum AIC information criterion. Once again, the log-likelihood statistic 171.5905(150.2075), AIC criterion -2.82(-2.7315), and the likelihood ratio test 42.7659 reject the linearity hypothesis of the stochastic process. The model reveals a significantly negative estimated coefficient on $VECM_{t-1}$ in the regime $s_t = 1$ (see Table 4.5) which may be interpreted as evidence of cointegration between the exchange rate and its fundamentals. The regime 2 reveals non-stationary behaviour denoted by a statistically insignificant negative coefficient on the $VECM_{t-1}$. The transition probabilities of the regimes are close, leading to durations that are quite similar between regimes. The lengths of deviation periods of the exchange rate from its fundamentals are quite similar to the length of long-run equilibrium periods. Moreover, the presence of explosive behaviour of the exchange rate is not observed. Hence, the data for France do not present evidence of bubbles driving exchange rate movements. Table 4.5 and Figure 4.6-France demonstrate the main computations and regime probabilities, respectively.

Table 4.5 – Main ML Estimation Results (France)

MSIAH(2)-VECM(4)*	Regime $s_t = 0$	Regime $s_t = 1$
Intercept	0.1671(5.3735)	0.1678(2.1981)
VECM$_{t-1}$	-0.049(-6.2178)	-0.035(-1.8672)
Standard Error	0.02191	0.05228
Duration	6.44 (50 obs.)	6(48 obs.)
Log-likelihood	171.5905 (150.2075)	
LR linearity test	42.765, $\chi^2(15) = [0.0002]^{**}$	
Transition Matrix	Regime $s_t = 0$	Regime $s_t = 1$
Regime $s_t = 1$	0.84	0.16
Regime $s_t = 2$	0.16	0.84

* Figures in parentheses denote t-statistics.

Figure 4.6-France



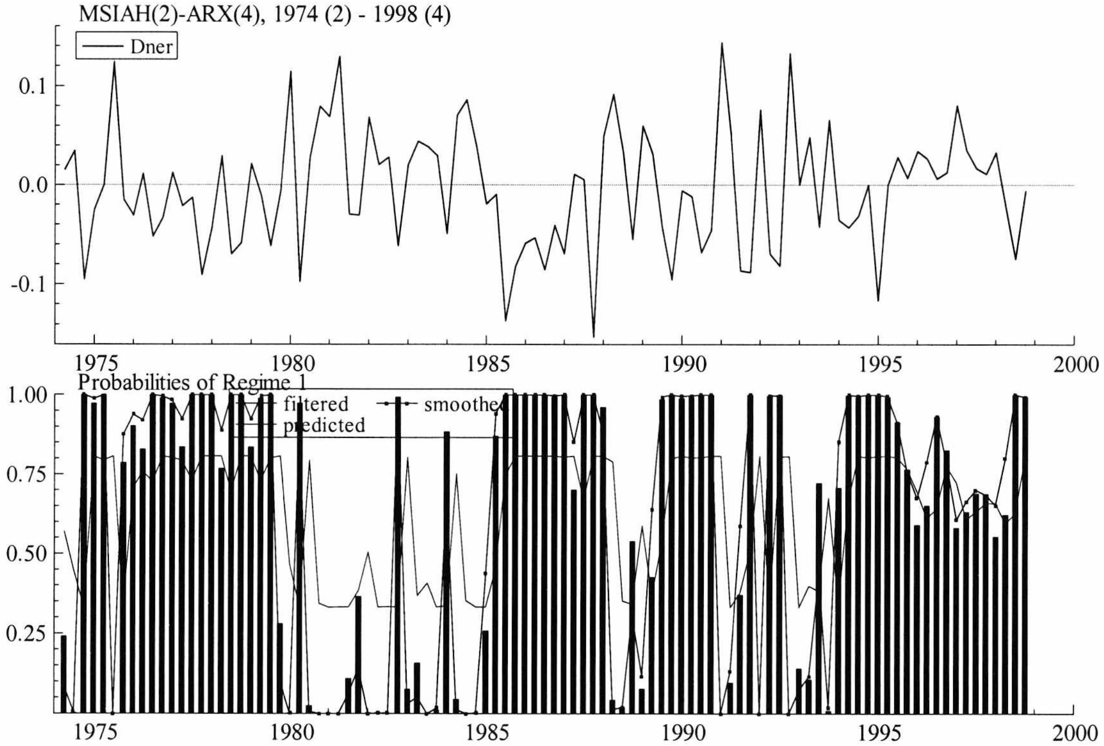
A MS(2)-VECM(4) is also selected for Germany based on the AIC information criterion. The log-likelihood statistic 161.0411(146.2912), AIC criterion -2.6069(-2.652) and likelihood ratio test 29.499 reject the linearity hypothesis of the system. The computations for coefficients on $VECM_{t-1}$ are found to be statistically insignificant which means that cointegration between the exchange rate and its fundamentals is not supported by the data. Deviations from the fundamentals solution are not corrected in the long run. Despite the coefficient on $VECM_{t-1}$ being positive in regime 1, indicating an explosive component in exchange rate behaviour, it is not significant. Furthermore, taking into account the duration for both regimes shows that the stochastic process defined by the non-stationary regime 0 prevails over the weak explosive path generated by the regime 1, which is reinforced by the high transition probability of leaving regime 1 to regime 0 (see Table 4.6). Thus, it is possible to conclude that bubbles are not driving the exchange rate movements in Germany for the sample period. Table 4.6 and Figure 4.7G display the main estimates and regime probabilities, respectively.

Table 4.6 – Main ML Estimation Results (Germany)

MSIAH(2)-VECM(4)*	Regime $s_t = 0$	Regime $s_t = 1$
Intercept	-0.0181(-0.5722)	0.0422(0.427)
$VECM_{t-1}$	-0.0269(-1.4633)	0.0089(0.0819)
Standard Error	0.03195	0.03376
Duration	5.18 (62 obs.)	3.01(37 obs.)
Log-likelihood	161.0411 (146.291)	
LR linearity test	29.499, $\chi^2(15) = [0.0139]^*$	
Transition Matrix	Regime $s_t = 0$	Regime $s_t = 1$
Regime $s_t = 1$	0.81	0.19
Regime $s_t = 2$	0.33	0.67

* Figures in parentheses denote *t*-statistics.

Figure 4.7-Germany



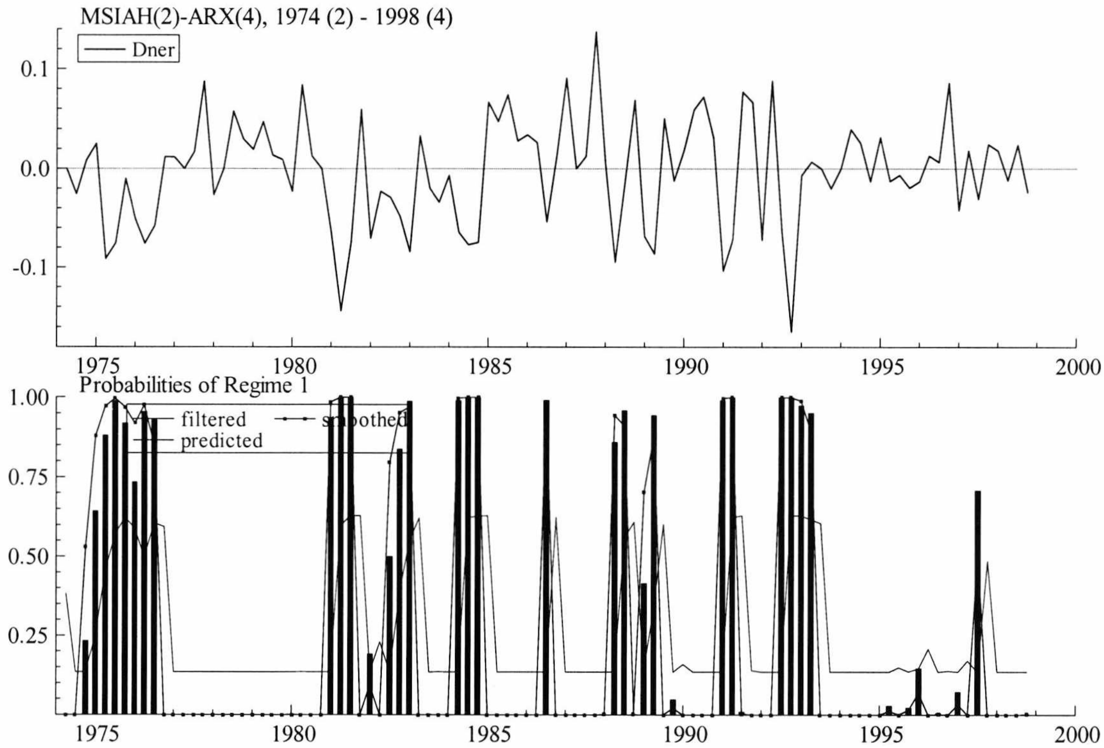
Finally, the results for the United Kingdom are estimated from a MS(2)-VECM(4) supported by the AIC information criterion. The log-likelihood statistic 178.239(157.811), AIC criterion -2.954(-2.885), likelihood ratio test statistic 40.855 all reject the linearity of the system. The estimates of the $VECM_{t-1}$ are also not significant, but both are negative, so that it is possible to infer that there is no cointegration between the exchange rate and its fundamentals (see Table 4.7). Furthermore, based on theses results the hypothesis of bubbles driving the exchange rate in the UK can be rejected. Table 4.7 and Figure 4.8UK display the main calculations.

Table 4.7 – Main ML Estimation Results (UK)

MSIAH(2)-VECM(4)*	Regime $s_t = 0$	Regime $s_t = 1$
Intercept	-0.2805(-2.433)	-0.0021(-0.0143)
VECM$_{t-1}$	-0.0517(-1.8213)	-0.006(-0.1643)
Standard Error	0.01474	0.03334
Duration	2.69 (26.9 obs.)	7.39(72 obs.)
Log-likelihood	178.2394 (157.8114)	
LR linearity test	40.8559, $\chi^2(15) = [0.0003]^{**}$	
Transition Matrix	Regime $s_t = 0$	Regime $s_t = 1$
Regime $s_t = 1$	0.63	0.37
Regime $s_t = 2$	0.13	0.87

* Figures in parentheses denote t-statistics.

Figure 4.8-UK



According to Psaradakis *et al.* (2004) the conventional tests of linear cointegration are reasonable in detecting long-run relationships. Even if occasionally the error process follows a non-stationary path due to different prevailing regimes, the tests are powerful enough to detect evidence of cointegration when the state indicators $\{s_t\}$ are uncorrelated. The conclusions found for cointegration in this section reinforce the results reached in the previous section using MS-unit root tests in rejecting the presence of bubbles in the exchange rates for these four industrialised market economies.

4.4 - Conclusion

The central idea of this chapter was to investigate the possibility of a bubble process driving the exchange rate for four industrialised economies The standard tests for unit roots and cointegration are unable to detect periodically collapsing bubbles. This type of stochastic process does not model regime shifts suitably and this affects test performances. The better option was to make use of an econometric approach which takes into account continuous regime shifts generated by bubbles collapsing periodically. To do so, a more flexible econometric technique is used, based on Markov switching (MS) regimes, where the parameters are defined conditional on an unobservable regime variable, and the transition between different regimes is determined by a homogeneous Markov chain.

The first econometric strategy consisted of applying the MS regime unit root test to two supposedly distinct regimes: one constitutes a fundamental solution and the other one constitutes a potential bubble solution. The tests were applied to Canada, France, Germany and the United Kingdom. The MS-regime unit root tests applied to each country did not reveal robustly the presence of bubbles governing the exchange rate, as the explosive roots detected in this variable were also detected in at least one of its fundamentals. The hypothesis of bubbles can only be supported fully if significant explosive roots are found in the exchange rate process, but not in its fundamentals. The bubble process is an independent stochastic process. Tests, nonetheless, detected non-linear behaviours for exchange rates.

Due to the potential weakness of MS-unit root tests a MS-ECM test was investigated based on a MS(M)-VECM(p) model with two distinct regimes (M=2), and different lag orders (*p*) for the VAR representation. The empirical evidence found in this analysis of the exchange rate in all four countries did confirm the results reached by the MS-regime unit root tests; namely that the interpretation of the estimates did not support the presence of bubbles. Although some countries have significant intercepts in different regimes, the estimated coefficients on the $VECM_{t-1}$ when positive, were not found to be significant. Tests also revealed non-linear behaviours for exchange rates.

Summarising, the hypothesis of periodically collapsing bubbles driving the exchange rate away from the fundamentals solution cannot be accepted for these four countries. Moreover, the Markov switching regime approach revealed significant non-linearities and different regimes. The existence of different regimes in the exchange rate is a finding that confirms previous results on non-linear exchange rate models [see Meese and Rose (1991), Taylor and Peel (2000), Yue and Kana (2000), Taylor and Peel and Sarno (2001), Kilian and Taylor (2003)]. However, these non-linearities do not appear to be linked to the explosive behaviour characteristic of processes driven by bubbles.

CHAPTER V

EFFECTS OF FOREIGN DEBTS

Introduction

In the previous two chapters exchange rate determination was analysed using the traditional monetary model of exchange rates. In chapter two, the monetary model was tested and the results pointed towards partial support for the monetary approach. In chapter three, the monetary model was used to investigate the presence of bubbles governing the exchange rate stochastic process. A MSECM[60] with two states was used: the fundamental solution state and the bubble solution state. The tests did not provide significant support for the hypothesis of bubbles driving the exchange rate away from its fundamentals. However, tests revealed some evidence of exchange rate nonlinearities.

For this chapter an alternative hypothesis based on financial fragilities is the focus of attention. The theoretical and empirical literature has demonstrated the importance of financial factors in explaining exchange rate movements; especially, financial variables that capture the foreign indebtedness of a country (see Devereux and Lane, 2003). Essentially, the literature highlights the effect of foreign debts, denominated in foreign currency, on exchange rate variability. The existence of un-hedged foreign currency denominated debts may bring about significant impacts on the exchange rate with potentially severe side effects on financial sectors and corporate balance sheets. A high level of un-hedged foreign currency denominated debts introduces additional economic costs to exchange rate variability. These adverse effects are expected to be more significant in emerging market economy

[60] Markov-switching error correction mechanism – MSECM.

countries than in developed market economy countries by virtue of credit constraints imposed by the international credit market.

In relation to international credit constraints, Eichengreen and Hausmann (1999) posit three different hypotheses on why some countries, especially developing countries, are credit-constrained and financially risky. The first one, called the **moral hazard hypothesis**, emphasises the distorting consequences of guarantees provided by governments to domestic financial market participants, and the possibility of international multilateral entities (IMF, for instance) rescuing countries in financial difficulties through emergency funds. It implies, initially, that foreign investors do not assume the full risk of their investment, creating an incentive to take excessively risky positions. However, extensive credit availability allows borrower countries to acquire excessive debts which generate future financial fragilities when credit constraints are imposed by the international financial community as a result of increasing default risks. Furthermore, some emerging economy countries which run a managed exchange rate regime make excessive un-hedged foreign currency borrowings. This distorts the financial flows towards the short term since foreign investors would be unwilling to assume long-term risky positions. These increasing un-hedged, short-term, foreign currency-denominated liabilities may be an additional source of exchange rate instability. Countries which adopt a more flexible exchange rate regime, on the other hand, are supposed to have lower levels of short-term capital flows leading to less volatile exchange rates. Flexible exchange rate regimes reflect better the economic fundamentals and increase the long-term credibility of their financial regime.

The second hypothesis, **the original sin hypothesis**, is related once again to the difficulties for some countries to borrow abroad or to borrow long term in domestic currency. This raises the problem of currency mismatches or maturity mismatches as long term domestic investment projects are often foreign currency-financed. If the exchange rate regime is flexible and governments allow for depreciation then currency mismatches take place[61]. On the other hand, if the exchange rate regime is managed and defended by monetary authority interventions, via rising interest rates, maturity mismatches take place (long-term investment projects will be financed by temporally high short-run interest rates). This leads the international financial market to constrain credit affecting exchange rate stability. Finally, the third hypothesis emphasises that financial constraints are also caused by **weakness of institutional and judicial infrastructures** that creates commitment and enforcement problems. Thus, Eichengreen and Hausmann (1999) conclude that many emerging market economies that own high levels of debt may have huge difficulties to manage their exchange rate volatility. These difficulties affect the financial credibility of a country leading to a more volatile exchange rate.

Devereux and Lane (2003) reinforce this argument and propose that high foreign debt stocks have differential impacts on exchange rate movements for developed market and emerging market countries. According to them, developed market countries have free access to the international financial market so that they can borrow by issuing assets denominated in their own currencies. Emerging market countries, in turn, are affected by severe borrowing constraints in international financial markets and only issue debts in foreign currencies. Hence an adverse external shock might produce perverse effects on the exchange rate in emerging economies. In particular, the imposition of credit constraints is a result of the increasing level of foreign indebtedness of developing economies

[61] Currency mismatches means, essentially, assets in domestic currency against liabilities in foreign currency.

leading to increasing default risks. A volatile and high risk premium usually arises in emerging market countries due to default risk in international credit operations, which can constrain foreign capital inflows[62]. These credit constraints associated with increased default risks can generate an additional capital outflow with destabilising impacts on exchange rate volatility in emerging economies. Hence, one can infer that the high stock of foreign debts has indirect effects on exchange rate volatility through its impact on the risk premium.

Some studies have also highlighted the importance of balance sheet effects[63] on macroeconomic outcomes. These effects may produce feedback impacts on exchange rate volatility and eventually on foreign indebtedness levels. Bernanke, Gertler and Gilchrist (1999) investigated the impact of balance sheet effects on business cycles. Extending this issue to the open economy is a recent important contribution to international finance and is reflected in the work of a number of authors; for example, Krugman (1999), Cook (2000), Aghion *et al.* (2001), Cespedes, Chang and Velasco (2001, 2004), Devereux and Lane (2001), and Eichengreen (2002). Cespedes *et.al.* (2004) investigate the dynamic impact of balance sheet effects on a financially vulnerable economy caused by high external indebtedness. According to them, in economies in financial distress the balance sheet effects are particularly magnified due to the increase in country risk with lasting and large adverse impacts on domestic economic variables. Actually, balance sheet effects may be a direct consequence of high levels of foreign currency-denominated debts and may generate a significant impact

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[62] Note that although a risk premium may be derived from different factors such as economic, social, political etc., in highly indebted economies economic factors leading to default risks may eventually prevail and reflect a significant percentage of the risk premium.

[63] Balance sheet effects are adverse consequences generated by financial variables shocks over firms' finances that have a high un-hedged level of debts. This causes an increase in default risks.

on exchange rate volatility. McKinnon and Pill (1998) analyse banking system efficiency and emphasise that in countries where the banking sector operates efficiently and is well regulated, the rapid credit expansion derived from a large foreign capital inflow is less likely to create risk conditions and to lead to financial and macroeconomic instability with adverse consequences for exchange rate volatility. By way of contrast, they state that in countries where credit institutions are not appropriately regulated and market failures exist in the financial system, bank credits are sometimes used for speculative purposes leading to an increase in systemic financial risks. Aghion *et al.* (2001) argue that, in credit-constrained economies with high levels of foreign currency-denominated debts, exchange rate volatility may generate adverse economic effects by deteriorating the balance sheets of private firms[64]. Pratap *et al.* (2003) investigated the role of currency mismatch in the Mexican corporate sector and found that exchange rate instabilities have adverse impacts on firms which hold dollar-denominated debts. Hausmann *et al.* (2001) state that monetary authorities, in turn, pay special attention to the exchange rate in countries in which there are large currency mismatches in the balance sheet of firms, banks, households or the government. According to them, if currency mismatches are large, exchange rate instabilities may bring about serious consequences to the private and public sectors, and even widespread bankruptcies in case of a large depreciation. Thus, difficulties arising from balance sheet problems may have impacts attenuated through the choice of exchange rate regimes.

Based on these arguments it is possible to conclude that there is a significant linkage between financial factors and exchange rate volatility. The introduction of balance sheet effects leads to an

[64] These effects are also strongly observed on government finances mainly in emerging market economies.

additional complicating ingredient to this analysis. The reasoning is that large stocks of un-hedged foreign currency-denominated debts generate exchange rate volatility with significant impacts on corporate balance sheets and additional increases in default risks. In fact, increases in costs of foreign currency-denominated debt repayments reduce corporate profits, the investment capacity, and thus the capital stock and output in the future. Ultimately, the conjunctions of these factors may lead to changes in market agents' expectations with potentially severe consequences for exchange rate volatility itself. In particular, these adverse consequences constitute a powerful fuel to the default risk via rises in economic uncertainties with feedback effects on the exchange rate variability. Additionally, and given the large impact that exchange rate volatility may have on highly indebted economies, monetary authorities will need to intervene in the foreign exchange market.

Devereux and Lane (2003) test the exchange rate vulnerability to external shocks in the presence of large stocks of un-hedged foreign currency denominated debts for a set of developed and developing countries. The hypothesis is that the existence of credit constraints, in combination with external debts, leads to a significant difference of optimal responses of exchange rates to shocks. They made use of a large cross-section of developed and developing countries to identify the main determinants of exchange rate volatility. Their results show that for developed countries bilateral exchange rate volatility is either positively affected by external financial linkages or affected insignificantly. In contrast, for developing countries financial linkages decrease sharply bilateral exchange rate volatility. Devereux and Lane also examine the importance of internal finance, capturing the degree of financial depth within countries, on exchange rate volatility and find that this variable increases exchange rate volatility for developed economies, but reduces it for developing countries. Hausmann *et al.* (2001) build measures of pass-through and of the ability of countries to borrow abroad in their own currency in order to explain exchange rate behaviour. They find a robust relationship between the ability of a country to borrow in its own currency and the way it manages its exchange rate policy. They also conclude that countries that do not borrow in their own currency hold higher levels of foreign reserves, and thus ensure much less exchange rate volatility supported by foreign reserves and interest rate interventions.

The research developed by Devereux and Lane (2003) explores the stochastic properties of cross section data to investigate the effect of debt variables on the bilateral exchange rate movements for a set of countries. However, the cross section analysis does not capture level variations over time. Essentially, the cross section modelling is a static analysis which investigates the effects of a change in one variable possibly between two time periods (20 year average growth rates for 100 countries for instance), but it does not allow isolating short run or long run impacts nor capturing the dynamics of adjustment. Also, a cross section assumes that the impact is the same for countries within the same category. Nevertheless, it is likely that considerable heterogeneity exists (financial sectors, supervision, credit constraints, etc.) across countries. In particular, cross sectional distributions demonstrate to be relatively stable and may hide a multitude of changes over time. Hence, cross section analysis is an econometric method that addresses part of the problem in examining variables that demonstrate this type of volatile behaviour. Alternatively, time series data modelling is an approach able to capture substantial changes over time when analysing effects of a variable on another. Essentially, it is able to identify, to measure and to predict effects of economic policies appropriately by relating variables behaviours at one point in time to others at another point in time. Furthermore, time series data modelling is an approach that allows for the researcher to work with a larger number of data points comparatively to the cross section data analysis. This property aggregates a relevant importance to investigations on financial variables as it increases considerably the degrees of freedom, reduces the collinearity among explanatory variables contributing to improve the efficiency of estimation techniques. Hence, the time-series data approach models more appropriately exchange rate series which are dynamic stochastic processes that show sustained periods of appreciation and depreciation over time.

Nonetheless, some variables in times series are not weakly exogenous so that some type of simultaneity relationship is a plausible hypothesis. This is the proposal of this chapter, as both exchange rate volatility and foreign indebtedness determine each other through the feedback effects commented earlier. Thus, a reasonable way of treating this difficulty is to assume a simultaneous behaviour for each variable. In this sense, it implies to let the time path of $\{y_t\}$ be affected by current and past realizations of $\{x_t\}$ and the time path of $\{x_t\}$ be affected by current and past realizations of $\{y_t\}$. This simple two variable system constitutes the central aspect of vector autoregression – VAR modelling – which introduces the idea of feedback effects between variables since the y_t and x_t sequences are allowed to affect each other.

This chapter explores the stochastic properties of time series data using VAR modelling in order to test, firstly, for the impact of the ratio of net foreign debt to GDP on exchange rate volatility and vice-versa. The key issue is that foreign debts may affect exchange rate volatility as result of increases in default risks, but feedback effects from impacts on corporate balance sheets may also take place and to affect the ratio of foreign debt to GDP[65]. The feedback effect constitutes the typical simultaneity and supports the basic properties inherently found in VAR modelling. The investigation is conducted by applying this econometric method to a sample of developed and emerging markets that show high and low levels of foreign indebtedness, respectively. The primary hypothesis tested is that changes in foreign debt impact on exchange rate volatility, but in a different way across developed and emerging market countries.

This type of modelling requires the use of structural shocks. According to Sims (1980), a structural VAR cannot be estimated directly due to feedback effects associated with the system. Hence, a Choleski decomposition is used in order to recover the structural residuals. The Choleski decomposition forces a potential asymmetry to the system by imposing an ordering to the variables. After the structural errors are recovered, the variance decomposition is produced to analyse the proportion of the movements in the exchange rate sequence due to its own shocks versus shocks to the foreign debt sequence. Finally, an impulse-response analysis is developed to trace out the time behaviour of a variable in response to various shocks (impulses) from another variable.

Additionally, according to Hausmann *et al.*(2001) even though emerging market countries are usually credit constrained to borrow abroad in their own currencies, their exchange rates may eventually reveal a low volatility by virtue of monetary authority interventions. Hausmann *et al.* report that the level of international

[65] Note also that firms, governments, and other market participants, tend to promote changes in their debt portfolio to prevent financial losses due to more volatile exchange rates.

reserves in countries under credit constraints and high un-hedged foreign currency indebtedness are, on average, higher and more volatile than countries with non-credit constraints. This may be a strategy which aims at controlling an eventual increase in exchange rate volatility. Moreover, the government authorities also tend to tighten monetary policy and to allow for more volatility in interest rates than in exchange rates. Calvo and Reinhart (2002) argue that the exchange rate in these countries seems to be floating with a "lifejacket" in reference to the high level of foreign reserves or interventions in interest rates. On the other hand, the exchange rate policy in developed market economy countries with non-credit constraints, and supposedly able to borrow in their own currency, may follow what Hausmann *et al.* called "benign neglect" to characterise limited interventions on the foreign exchange market and the feasibility for the exchange rate to be more volatile. The impact of exchange rate volatility on interest rates and international reserves can be viewed as a class of spillover effects (see Hong, 2001).

Table 5.1 reports some basic statistics for a four-country sample made up of Australia, Canada, Mexico and Turkey and used in this study. The interest rate and international reserves are variables that may reflect some type of spillover effects as suggested by Hausmann *et al.*(2001) and Calvo and Reinhart (2002). The computations reported in Table I show evidences that high levels of foreign debts may have stronger impacts on the exchange rate volatility in financially fragile economies. Note that, the ratio of net foreign debt to GDP and the exchange rate volatility, as well as interest rate volatility, observed for Mexico and Turkey are, on average, substantially higher than for Australia and Canada. These observations reinforce the relevance of the basic hypothesis investigated in this study.

Table 5.1 – Basic Statistics by Country*'**

Country	Average				Variance		
	ner	debt	ir	Res	debt	ir	res
Australia	0.063	17.95	9.42	2.78	0.0012	17.292	0.0006
Canada	0.024	31.11	8.26	2.35	0.0224	16.655	0.0001
Mexico	0.942	42.89	39.86	4.61	0.0223	677.51	0.0005
Turkey	$6.44e^{12}$	46.57	67.99	6.91	0.0461	1009.1	0.0003

* The statistics are based on 1980:1–2002:1 period, except for Turkey (1987:1-2002:1).
** ner: quarterly nominal exchange rate variance, debt: foreign debt/gdp in %, ir: interest rate in %, res: international reserves/gdp in %.

This chapter also investigates the role of monetary authorities in response to exchange rate volatility. The basic hypothesis tests if exchange rate volatility is followed by increases in either interest rate or international reserves volatility. Given that the interest rate and international reserves are assumed to be stabilising tools, an additional vector is introduced into the VAR model in order to investigate if exchange rate volatility due to high foreign debts affects interest rate and international reserves. This strategy aims at capturing eventual spillover effects caused by foreign debts shocks on either the interest rate or international reserves.

The chapter is structured in five distinct sections. Section 5.1 outlines the theoretical considerations to support the underlying ideas. Section 5.3 describes the statistical data, variables used and collecting sources. Section 5.4 presents the empirical results with a brief explanation of the methodology adopted. Finally, Section 5.5 presents the conclusions.

5.1 - Theoretical Considerations

The increasing level of foreign indebtedness and its macroeconomic effects has assumed considerable importance in academic studies in the last two decades. Some theses studies, as cited in the introduction, describe real and nominal impacts, resulting from uncontrolled growth of external debts, on the economy of different countries. This chapter aims at giving an additional contribution to the debate by focusing on the effects of foreign debts on exchange rate volatility (nominal shocks). The theoretical background follows ideas derived from Bernanke, Gertler and Gilchrist (1999), Obstfeld and Rogoff (1995, 2000), Devereux and Lane (2003), and considerations about risk premiums in exchange rates from Sarno and Taylor (2002). In essence, the main objective is to associate the increasing foreign indebtedness to default risks in conjunction with credit constraints, its impact on exchange rate volatility generated by two types of effects: (1) direct effects, and (2) feedback effects.

The underlying idea is that the higher the level of un-hedged foreign debts the higher the default risk. So a higher level of foreign debts implies an implicit default risk which contaminates the market. This increasing risk raises the risk premium due to a generalised default fear in the market resulting in the exchange rate diverging from its expected change determined by fundamentals[66]. Thus, a high probability of default may interfere with market agent's expectations and cause temporary turbulences in exchange rate volatility due to a higher volatility in risk. This constitutes the direct effect of high stocks of foreign debt on exchange rate volatility. The direct effect also causes credit constraints as a result

[66] Note that the risk percentage embodied into the interest rate leads to an additional increase in default risk as corporate financial costs are affected. In practice, the risky economic context determines simply credit constraints to highly indebted economies.

of increases in default risk due to excessive levels of borrowings and may have additional side effects on exchange rate volatility. Feedback effects, in turn, denote the potential impact of the exchange rate volatility on corporate balance sheets, consequently on the ratio of foreign debts to GDP. Potentially, firms tend to change their debt portfolios from foreign sources to domestic sources over time as it may be costly to keep foreign currency liabilities in a context of volatile exchange rate. The theoretical framework for feedback effects is not presented in this chapter as it does not constitute the central topic of this chapter.

The theoretical argument of direct effects describes the relationship between credit constraints and risk premiums based on a single small economy made up of households consuming one imported good, and supplying labour. This hypothetical economy also produces a non-traded good and an export good. Apart from the non-traded good, prices of both import and export goods are set in the international market. Firms in the export sector use both labour and intermediate imports and do not have any power over international prices. Firms in the intermediate import sector acquire inputs abroad and transfer them to exporters. The problem arises when firms have credit constraints for purchases of intermediate inputs and need to borrow an amount B[67] from foreign banks to finance their purchases. If there are default risks, the intermediate input importers are required to pay a fee per unit of imported input. This additional charge is a risk premium linked to the default risk and it is supposed to increase according to the amount borrowed in relation to net worth. Increasing the un-hedged amount of B may lead to corporate financial fragilities and to rising risk premiums. Converted into the domestic price by a nominal exchange rate, the

[67] International exchange is usually made via financings in foreign currency, rather than in cash, from international credit markets.

variable cost (VC) of the input for the importer is expressed as follows:

$$\text{VC} = sp^{*}q[e^{z(\alpha)}] \qquad (5.1.1)$$

where s is the nominal exchange rate, p^* is the imported input price, and q is the quantity imported. The term $z(\alpha)$ denotes an additional risk premium function associated with the importer's default risk.

The risk premium rate (α) depends on the ratio of net corporate balance sheet to P_x, $\varphi = \frac{1}{sP_x}[sB - A]$, in which P_x is the traded good price, B and A are the borrowings in foreign currency and assets in domestic currency, respectively, so that $\alpha = f(\varphi)$[68]. If $\varphi > 0$ implies un-hedged foreign debts and potential risk of default.

By assuming in this context that q is obtained via B, the expression (5.1.1) in conjunction with φ indicate that an adverse exchange rate shock may inflate corporate liabilities, government finances[69], and lead to balance sheet effects if indebted firms/governments do not have an appropriate coverage in assets[70]. This is the essence of direct effects from high levels of foreign debts on exchange rate volatility caused by corporate balance sheet effects.

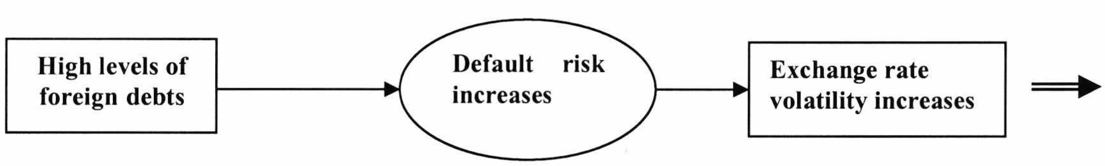
The linkage between the default risk and the risk premium leads to collateral effects on exchange rate volatility. If

[68] For simplicity, it is assumed that the default risk corresponds to 100% of the risk premium.
[69] Government finances are affected when loans to the private sector are guaranteed by the public sector (moral hazard hypothesis).
[70] This situation may also compromise future corporate and government investment abilities with impacts on eventual business cycles [see Bernanke, Gertler and Gilchrist (1999)].

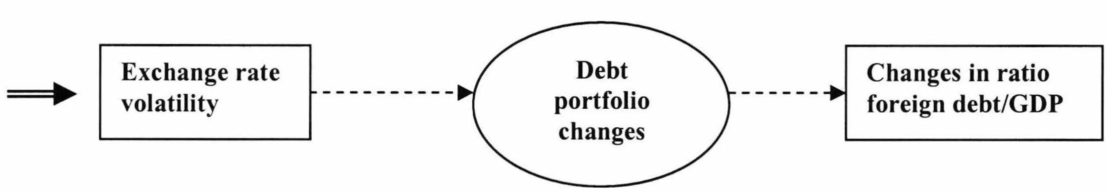
the function $z(\alpha)$ in (5.1.1) satisfies $z'(\alpha)>0$, then corporate balance sheet positions assume a significant importance of affecting exchange rate volatility through having implications on the default risk and the risk premium in exchange rates. The expression $z'(\alpha)>0$ constitutes an important rationale to understand the impact of high stock of foreign debts on exchange rate volatility via default risks (risk premiums) and feedback effects causing balance sheet changes. Figure 5.1 describes the two underlying general hypotheses assumed in this context.

Figure 5.1 – Direct Effects and Feedback Effects*

1- Direct Effects



2- Feedback Effects



*** Full line arrows are direct effects and dashed line arrows are feedback effects. Note, additionally, that direct effects from high stocks of foreign debts on exchange rate volatility may potentially produce feedback effects on foreign debts/GDP as debt portfolio changes take place. Firms will tend to shift their businesses to domestic credits in presence of lasting exchange rate shocks.**

Note that a risk component does not make part of the set of exchange rate fundamentals. In fact, it is an extraneous element that affects the exchange rate behaviour and violates a crucial assumption of the efficient foreign exchange market hypothesis: the risk-neutral assumption. Indeed, the default risk is embodied in the risk premium, and under temporary circumstances has side effects on the exchange rate volatility[71]. Hence, if the market participants, especially in foreign exchange markets, are risk-averse, the uncovered interest parity (UIRP) condition based solely in economic fundamentals, and expressed by $i_t - i_t^* = \Delta_k^e s_{t+k}$, does not hold due to the distortion created by the risk premium. Agents will require a higher rate of return than the simple interest rate differential in return for the exchange rate risk assumed. The following expression reflects the effects of increased risks [$z'(\alpha) > 0$] relative to the possibility of defaults:

$$-(i_t - i_t^*) = \Delta_t^e s_{t+1} + \alpha_t \qquad (5.1.2)$$

where i_t and i_t^* denote the domestic and foreign interest rate, $\Delta_t^e s_{t+1}$ the expected change in the exchange rate and α_t is the risk premium.

The empirical literature points out that the existence of a risk premium has important implications for exchange rate regressions. Hence, if the efficient foreign exchange market hypothesis holds, then the UIRP condition can be tested as follows:

[71] In this context, temporary circumstances refer to adverse shocks that affect the international confidence in credit markets.

$$\Delta_k^e s_{t+k} = \alpha_t + \beta\left(f_t^k - s_t\right) + v_{t+k} \qquad (5.1.3)$$

where f_t^k is the forward exchange rate at time k, α_t is the risk premium which must be equal to zero for UIRP to hold and v_{t+k} is white noise error at time $t+k$.

The term $f_t^k - s_t$ in (5.1.3) denotes the forward premium predicted in a forward exchange rate contract[72]. If the information set Ω_t available to the market contains α_t, and the rational expectations hypothesis holds, then the forward premium should be enough to cover the expected change in the exchange rate and the risk premium component.

Thus, when $\alpha = 0$ the estimated value of the slope parameter β is expected to be equal to one and the disturbance term v_{t+k} to be uncorrelated with the information available at time t. Fama (1984), nevertheless, derived a prior version of the estimator for β in (5.1.3) that did not take into account rational expectations as:

$$\beta = \frac{\operatorname{cov}\left(\Delta s_{t+k}, f_t^k - s_t\right)}{\operatorname{var}\left(f_t^k - s_t\right)} \qquad (5.1.4)$$

By applying rational expectations to a version of Fama (1984) and assuming $\alpha \neq 0$, Sarno and Taylor (2002) suggest that the

[72] The basic difference between UIRP and covered interest rate parity (CIRP) is that the latter concept implies the coverage of a forward exchange rate contract.

estimator for the coefficient β in (5.1.3) can be estimated via OLS as follows[73]:

$$\beta = \frac{\operatorname{var}\left(\Delta_k^e s_{t+k}\right) + \operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right)}{\operatorname{var}\left(\alpha_t\right) + \operatorname{var}\left(\Delta_k^e s_{t+k}\right) + 2\operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right)} \qquad (5.1.5)^{74}$$

where cov and var are the covariance and variance, respectively.

A considerable number of empirical studies has reported a negative β in (5.1.3) and particularly close to minus unity [see Froot and Thaler (1990) and Sarno and Taylor (2002)][75]. These unexpected findings appear to violate the UIRP condition as they indicate that the exchange rate should appreciate the more the domestic interest rate exceeds the foreign interest rate. However, the rejection of UIRP based on a negative β may be a misleading interpretation because it does not usually take into account the effect of the risk premium (α_t) associated with an excess supply of the domestic currency (see Hodrick, 1986, 1992). Essentially, the UIRP simply states that the exchange rate is expected to depreciate at $t+k$ when a risk premium is present resulting in the $\operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right) < 0$ in (5.1.4). The negativity of β is compensated by rises in $\operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right) < 0$ due to increasing holdings of the domestic currency by virtue of the interest rate differential. In reality, a positive

---

[73] Note that $\operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right) < 0$ implies effects of the risk premium (α_t) on the expected change in exchange rate ($\Delta_k^e s_{t+k}$), and vice-versa, so that α_t and β in (5.1.3) follow opposite stochastic paths.

[74] From (5.1.3) and (5.1.4), it can implicitly be inferred that $(f_t^k - s_t) = \Delta_k^e s_{t+k} + \alpha_t$ or the forward premium $FP = \Delta_k^e s_{t+k} + \alpha_t$ so that $var(PF) = var(\Delta_k^e s_{t+k} + \alpha_t)$. As $\Delta_k^e s_{t+k}$ *and* α_t are not independent, so $var(PF) = var(\alpha_t) + var(\Delta_k^e s_{t+k}) + 2\,cov(\Delta_k^e s_{t+k}, \alpha_t)$ as in (5.1.5).

[75] The literature usually denominates the negativity of β in (5.1.3) as a forward discount bias.

interest rate differential in (5.1.2) suggests that the foreign currency is at a premium in the forward market which implies a time-varying and increasing risk premium. As a result, the longer and the larger a positive interest rate differential remains the more the risk component (α_t) becomes volatile over time, so that the efficient market hypothesis is rejected.

To demonstrate this, Sarno and Taylor (2002) suggest taking a weaker proposition for the coefficient β in (5.1.5) so that it is assumed to be less than ½. This assumption implies a significant risk premium ($\alpha \neq 0 \Rightarrow \beta \neq 1$) included in (5.1.3)[76], and the estimator for β can be rearranged as follows:

$$\beta = \frac{\operatorname{var}\left(\Delta_k^e s_{t+k}\right) + \operatorname{cov}\left(\alpha_t, \Delta_i^e s_{t+k}\right)}{\operatorname{var}\left(\alpha_t\right) + \operatorname{var}\left(\Delta_i^e s_{t+k}\right) + 2\operatorname{cov}\left(\alpha_t, \Delta_i^e s_{t+k}\right)} < \frac{1}{2} \qquad (5.1.6)$$

The expression (5.1.6) shows that the estimation of β is affected by the effect of the risk premium component on the expected change in the exchange rate denoted by $\operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right) \neq 0$ so that:

$$2\operatorname{var}\left(\Delta_k^e s_{t+k}\right) + 2\operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right) < \operatorname{var}\left(\alpha_t\right) + \operatorname{var}\left(\Delta_k^e s_{t+k}\right) + 2\operatorname{cov}\left(\alpha_t, \Delta_k^e s_{t+k}\right)$$

which results in:

$$\operatorname{var}\left(\alpha_t\right) > \operatorname{var}\left(\Delta_k^e s_{t+k}\right) \qquad (5.1.7)$$

[76] In the context of foreign debts effects on the exchange rate volatility, a risk premium is given rise in the market as a result of default risks. The existence of a risk premium implies an inefficient foreign exchange market.

The inequality (5.1.7) states that the variance of the risk premium is greater than the variance of the expected change in the nominal exchange rate[77]. Equivalently, this expression also states that the existence of risk premiums leads to a larger volatility in the exchange rate than the volatility determined solely by economic fundamentals[78]. In conclusion, excessive levels of foreign indebtedness lead to the risk premium component increasing, constraining credit inflows and causing the exchange rate to be more volatile.

The above arguments allow deriving the main hypothesis investigated in this chapter. For industrialised market economies, which are not subject to credit constraints in the international financial market and the default risk is almost nonexistent, foreign debts are expected to have low impacts on the variance of the exchange rate. However, for emerging market economies which have usually a high level of foreign indebtedness, huge credit constraints associated to default risks, foreign debts are expected to cause significant impact on the variance of the exchange rate. As a corollary from these arguments, the hypothesis of monetary policy interventions to minimise eventual destabilising effects of foreign debts on the exchange rate variation is also tested.

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[77] Cespedes *et al.* (2004) developed a theoretical model in which they show that exchange rate shocks raise the country risk premium in financially vulnerable economies. This finding reinforces the argument that an increase in the risk premium may lead to more volatile exchange rates.

[78] The forward exchange rate is determined by agents' future expectations about economic fundamentals plus a volatile risk premium. The introduction of a risk premium constitutes a disturbance and a violation of the efficient market hypothesis as expressed in (2.3) by $\alpha \neq 0$.

5.2 – Data

The data are collected from the International Financial Statistic-IFS provided by the International Monetary Fund for quarterly data, and from the World Bank Database for yearly data and from the Central Bank of each country for the daily nominal exchange rate. They consist of time series of two developed market countries (Australia and Canada) and two emerging market countries (Mexico and Turkey). The period of analysis extends from 1980:1 to 2002:1. The exception is Turkey for which the period of analysis extends from 1987:1 to 2002:1 due to data availability.

This study conducts the analysis primarily by using the quarter by quarter daily variance of the nominal exchange rate (ner) against the US dollar to denote the exchange rate volatility, the ratio of net foreign debt to GDP (net foreign debt = gross foreign debt – international reserves)[79], the ratio of international reserve to GDP and the interest rate country by country. The quarterly exchange rate volatility is the variance of its daily close value quarter by quarter. The process of estimation of the net foreign debt consists, initially, of calculating a proxy for the gross foreign debt by using the methodology proposed by Lane and Ferretti (1999). They show a practical methodology to construct estimates of foreign assets and liabilities, equity and debt subcomponents for 66 industrial and developing countries. Based on Lane and Ferretti (1999), the estimation for the gross foreign debt consists firstly of summing the portfolio debt liabilities to borrowings from the IMF and to exceptional financings[80] collected from the balance of

[79] In essence, the concept of net foreign debts is used as a proxy to $\left[sB-A\right]$ as described in the previous section.

[80] Exceptional financing includes arrears on payment of principal and interest on foreign liabilities, loans contracted for balance of payment needs, debt reductions or forgiveness of debts. Although exceptional

payments data available in the IFS/IMF. Next, this flow value is summed to the foreign debt's stock value of the previous year provided by the World Bank database and so successively year on year. Finally, the net foreign debt is obtained by subtracting the international reserves balance from the gross foreign debt estimates calculated previously. The interest rate is the money market rate which is supposed to reflect better the liquidity conditions in the money market. The international reserves data and GDP are provided by IFS/IMF.

Section 4 makes use of VAR modelling to test for the effects of foreign debts on exchange rate volatility and eventual monetary interventions in the foreign exchange market.

5.3 – Empirical Results

As mentioned earlier, the high level of foreign debt has become an important variable in macroeconomic debates and may have an important role in explaining the huge volatility in exchange rates. VAR modelling has been used extensively in the empirical macroeconomics literature by assuming simultaneity between variables[81]. This approach is applied to this chapter to investigate the ability of changes in foreign debts to explain exchange rate volatility and vice-verse[82].

It is worth emphasising that the main objective of this investigation is to analyse the economic significance of the interrelationships among variables under analysis in the study, and

financing is recorded below the line in the balance of payment as a financing item, it is assumed as external liabilities in order to have a more accurate figure.

[81] See the seminal work of Sims (1980).

[82] Enders (2004) states that it is convenient to use the same lag length for each variable in order to preserve the symmetry of the system and to obtain efficient estimates by OLS.

not simply to select models for forecasting purposes. This means that data stationarity is an important requirement in implementing this type of investigation. Even though some information concerning long-run relationships is wasted, a suitable differencing process may be needed in order to meet the stationarity required. The usual ADF test is a traditional tool which helps to identify the integration order of the variables. For simplicity, the maximum number of lags for the ADF test is fixed to three[83].

The results are displayed in Tables 5.2 and 5.3 as follows:

Table 5.2 – Unit Root Tests

	Variables in Level							
	Vner		Ndebt/gdp		IR		res/gdp	
Country	L	t-adf	L	t-adf	L	t-adf	L	t-adf
Aust	0	-15.6**	0	-1.34	0	-1.30	0	-2.54
	1	-10.67**	1	-1.27	1	-1.74	1	-1.91
	2	-8.65**	2	-1.35	2	-2.17	2	-1.87
	3	-7.93**	3	-1.44	3	-1.83	3	-1.89
Can	0	-6.28**	0	-1.37	0	-1.89	0	-0.93
	1	-6.06**	1	-1.29	1	-1.58	1	-0.64
	2	-4.32**	2	-1.209	2	-1.39	2	-0.52
	3	-3.54**	3	-1.204	3	-1.79	3	-0.60
Mex	0	-7.62**	0	-1.20	0	-2.25	0	-2.10
	1	-4.97**	1	-1.15	1	-1.95	1	-2.11
	2	-3.79**	2	-1.20	2	-1.40	2	-2.36
	3	-3.73**	3	-1.15	3	-1.49	3	-2.31
Turkey	0	-5.96**	0	-2.70	0	-4.25**	0	-4.25**
	1	-4.47**	1	-1.25	1	-4.93**	1	-3.14*
	2	-3.68**	2	-1.26	2	-3.36*	2	-3.36*
	3	-3.07*	3	-1.30	3	-3.41*	3	-3.41*

**** 1% level of significance and * 5% level of significance.**

[83] For the ADF test, the autoregressive lag k is selected from the range $0 < k \leq 8$ plus a constant term by using a sequential t-test (at 5% level) for the significance of the coefficient on the longest lag. The tests revealed that k=3 is, on average, the most appropriate.

Table 5.3 – Unit Root Tests

	Variables in First Difference							
	Vner		Ndebt/gdp		IR		res/gdp	
Country	**L**	**t-adf**	**L**	**t-adf**	**L**	**t-adf**	**L**	**t-adf**
Aust	0	stationary	0	-9.02**	0	-7.41**	0	-11.04**
	1		1	-5.52**	1	-5.27**	1	-6.86**
	2		2	-4.39**	2	-4.96**	2	-5.33**
	3		3	-3.72**	3	-4.68**	3	-4.08**
Can	0	stationary	0	-6.73**	0	-9.68**	0	-10.16**
	1		1	-4.23**	1	-7.10**	1	-6.93**
	2		2	-3.03*	2	-4.31**	2	-4.77**
	3		3	-2.85	3	-4.28**	3	-3.79**
Mex	0	stationary	0	-8.36**	0	-9.45**	0	-5.84**
	1		1	-5.65**	1	-8.22**	1	-3.83**
	2		2	-4.76**	2	-5.55**	2	-4.87**
	3		3	-3.73**	3	-4.75**	3	-4.83**
Turkey	0	stationary	0	-11.04**	0	stationary	0	stationary
	1		1	-6.08**	1		1	
	2		2	-4.65**	2		2	
	3		3	-4.21**	3		3	

** 1% level of significance and * 5% level of significance.

The results demonstrate that the exchange rate variance (volatility) is stationary in levels for all countries in this sample. The variables Ndebt/gdp, interest rate (IR) and res/gdp are integrated of order one, I(1), for all countries except Turkey. For Turkey the interest rate (IR) and res/gdp are I(0). So the empirical analysis uses the exchange rate variance, the interest rate and res/gdp in levels for Turkey. The other variables are in first differences.

As a VAR model, in general, generates a high loss of degrees of freedom the more variables are included in the model, it is recommended to fix parsimoniously the lag length. The empirical literature suggests a typical procedure for data in quarterly periodicity which consists of beginning from a model with 12 lags and gradually reducing the lag order by applying a likelihood ratio test[84]. According to Enders (2004), this test has an asymptotic χ^2 distribution with degrees of freedom equal to the number of restrictions in the system. Given a suitable significance level, the null hypothesis of a lower lag length is accepted if the calculated value is less then the critical χ^2 statistic. The tests are applied firstly to a two-vector VAR (Vner x Ndebet/gdp) and next applied to a three-vector VAR (Vner x Ndebt/gdp x IR or res/gdp). Although not reported, the results reveal that a 6th order VAR dominates a 12th order VAR and a 3rd order VAR dominates a 6th order VAR for all countries without a significant loss of explanatory power. Hence, the models estimated in this study follow a 3rd order VAR[85].

Two structural models, and their respective reduced forms, are investigated. The first one, structural model I, is a two vector VAR made up of exchange rate volatility (V_t) and the ratio of net foreign debt to GDP (D_t). The second one, structural model II, is a three vector VAR made up of exchange rate volatility, the ratio of net foreign debt to GDP and the interest rate or international reserves (Z_t).

The structural Model I is as follows:

$$\begin{aligned} V_t &= \alpha_{10} - \phi_{12}(0)D_t + \sum_{k=1}^{3}\beta_{11}(k)V_{t-k} + \sum_{k=1}^{3}\phi_{12}(k)D_{t-k} + \varepsilon_{vt} \\ D_t &= \alpha_{20} - \beta_{21}(0)V_t + \sum_{k=1}^{3}\beta_{21}(k)V_{t-k} + \sum_{k=1}^{3}\phi_{22}(k)D_{t-k} + \varepsilon_{dt} \end{aligned} \qquad (5.3.1)$$

---

84 The likelihood ratio test is defined by the following expression: $(T-c)\left(\log\left|\Sigma_r\right|-\left|\Sigma_u\right|\right)$ where T = number of observations , c = number of parameters estimated in each equation of the restricted system, and $\log\left|\Sigma_n\right|$= natural log of the determinant of the variance/covariance matrix of the restricted and unrestricted models, respectively.

85 Note that dummy variables have also been included in the empirical model to capture the effect of outliers.

And its matrix form is defined as:

$$\begin{bmatrix} 1 & \phi_{12}(0) \\ \beta_{21}(0) & 1 \end{bmatrix} \begin{bmatrix} V_t \\ D_t \end{bmatrix} = \begin{bmatrix} \alpha_{10} \\ \alpha_{20} \end{bmatrix} + \begin{bmatrix} \sum\limits_{k=1}^{3} \beta_{11}(k) & \sum\limits_{k=1}^{3} \phi_{12}(k) \\ \sum\limits_{k=1}^{3} \beta_{21}(k) & \sum\limits_{k=1}^{3} \phi_{22}(k) \end{bmatrix} \begin{bmatrix} V_{t-k} \\ D_{t-k} \end{bmatrix} + \begin{bmatrix} \varepsilon_{vt} \\ \varepsilon_{dt} \end{bmatrix} \qquad (5.3.2)$$

If $H_1 = \begin{bmatrix} 1 & \phi_{12}(0) \\ \beta_{21}(0) & 1 \end{bmatrix}$, $X_t = \begin{bmatrix} V_t \\ D_t \end{bmatrix}$, $\Psi_0 = \begin{bmatrix} \alpha_{10} \\ \alpha_{20} \end{bmatrix}$, $\Psi_k = \begin{bmatrix} \sum\limits_{k=1}^{3} \beta_{11}(k) & \sum\limits_{k=1}^{3} \phi_{12}(k) \\ \sum\limits_{k=1}^{3} \beta_{21}(k) & \sum\limits_{k=1}^{3} \phi_{22}(k) \end{bmatrix}$,

$\varepsilon_{it} = \begin{bmatrix} \varepsilon_{vt} \\ \varepsilon_{dt} \end{bmatrix}$, then,

$$X_t = A_0 + \sum_{k=1}^{3} A_k X_{t-k} + e_t \qquad k = 1,2,3. \qquad (5.3.3)$$

where,

$$A_0 = H^{-1}\Psi_0, \; A_k = H^{-1}\Psi_k, \; e_t = H^{-1}\varepsilon_t$$

From (5.3.3) the reduced form model I can be decomposed into two single equations:

$$\begin{aligned} V_t &= a_{10} + \sum_{k=1}^{3} a_{11}(k) V_{t-k} + \sum_{k=1}^{3} a_{12}(k) D_{t-k} + e_{1t} \\ D_t &= a_{20} + \sum_{k=1}^{3} a_{21}(k) V_{t-k} + \sum_{k=1}^{3} a_{22}(k) D_{t-k} + e_{2t} \end{aligned} \qquad (5.3.4)$$

where,

$$\begin{aligned} e_{1t} &= \left[\varepsilon_{vt} - \phi_{12}(0)\varepsilon_{dt}\right] / \left[1 - \beta_{21}(0)\phi_{12}(0)\right] \\ e_{2t} &= \left[\varepsilon_{dt} - \beta_{21}(0)\varepsilon_{vt}\right] / \left[1 - \beta_{21}(0)\phi_{12}(0)\right] \end{aligned} \qquad (5.3.5)$$

The structural model II can also be derived as:

$$V_t = \alpha_{10} - \phi_{12}(0)D_t - \phi_{13}(0)Z_t + \sum_{k=1}^{3}\beta_{11}(k)V_{t-k} + \sum_{k=1}^{3}\phi_{12}(k)D_{t-k} + \sum_{k=1}^{3}\gamma_{13}(k)Z_{t-k} + \varepsilon_{vt}$$

$$D_t = \alpha_{20} - \phi_{21}(0)V_t - \phi_{23}(0)Z_t + \sum_{k=1}^{3}\beta_{21}(k)V_{t-k} + \sum_{k=1}^{3}\phi_{22}(k)D_{t-k} + \sum_{k=1}^{3}\phi_{23}(k)Z_{t-k} + \varepsilon_{dt} \quad (5.3.6)$$

$$Z_t = \alpha_{30} - \phi_{31}(0)V_t - \phi_{32}(0)D_t + \sum_{k=1}^{3}\beta_{31}(k)V_{t-k} + \sum_{k=1}^{3}\phi_{32}(k)D_{t-k} + \sum_{k=1}^{3}\phi_{33}(k)Z_{t-k} + \varepsilon_{zt}$$

And following the same steps as before, the following representation is obtained:

$$Y_t = B_0 + B_k \sum_{k=1}^{3} Y_{t-k} + e_t \qquad k = 1,2,3. \qquad (5.3.7)$$

where,

$$B_0 = H^{-1}\Psi_0,\ B_k = H^{-1}\Psi_k,\ e_t = H^{-1}\varepsilon_t$$

The reduced-form model II is then derived from (5.3.7):

$$V_t = b_{10} + \sum_{k=1}^{3} b_{11}(k)V_{t-k} + \sum_{k=1}^{3} b_{12}(k)D_{t-k} + \sum_{k=1}^{3} b_{13}(k)Z_{t-k} + e_{1t}$$

$$D_t = b_{20} + \sum_{k=1}^{3} b_{21}(k)V_{t-k} + \sum_{k=1}^{3} b_{22}(k)D_{t-k} + \sum_{k=1}^{3} b_{23}(k)Z_{t-k} + e_{2t} \qquad (5.3.8)$$

$$Z_t = b_{30} + \sum_{k=1}^{3} b_{31}(k)V_{t-k} + \sum_{k=1}^{3} b_{32}(k)D_{t-k} + \sum_{k=1}^{3} b_{33}(k)Z_{t-k} + e_{3t}$$

where,

$$
\begin{aligned}
e_{1t} &= \left[\varepsilon_{vt} - \phi_{12}(0)\varepsilon_{dt} - \gamma_{13}(0)\varepsilon_{zt}\right]/\Delta \\
e_{2t} &= \left[\varepsilon_{dt} - \beta_{21}(0)\varepsilon_{vt} - \gamma_{23}(0)\varepsilon_{zt}\right]/\Delta \qquad (5.3.9) \\
e_{3t} &= \left[\varepsilon_{zt} - \beta_{31}(0)\varepsilon_{vt} - \phi_{32}(0)\varepsilon_{dt}\right]/\Delta
\end{aligned}
$$

$$
\Delta = \left[1 + \phi_{12}(0)\gamma_{23}(0)\beta_{31}(0) + \beta_{21}(0)\phi_{32}(0)\gamma_{33}(0) - \beta_{31}(0)\gamma_{13}(0) - \beta_{21}(0)\phi_{12}(0) - \phi_{32}(0)\gamma_{23}(0)\right]
$$

From (5.3.5) and (5.3.9), it is possible to conclude that the error terms e_{1t}, e_{2t} and e_{3t} are composites of the underlying primitive errors ε_{vt}, ε_{dt} and ε_{zt}, respectively. The forecast error terms produced by the reduced-form model need to be decomposed into their structural forms to allow an appropriate economic interpretation. Although the traditional literature states that Choleski decomposition is devoid, in general, of any particular economic interpretation by imposing random restrictions (see Enders, 2004), it, in fact, forces a potential and important asymmetry to the system that may have theoretical foundations[86]. Essentially, the mechanics inherent to this method allows recovering the structural shocks ε_{vt}, ε_{dt} and ε_{zt} from e_{1t}, e_{2t} and e_{3t} with the economic sense determined by the own researcher.

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[86] Although other methods of errors decomposition could have been considered, the Cholesky decomposition is essentially simpler, allows for some economic rationale and meets the basic objectives of this investigation. Beveridge and Nelson (1981) decomposition, for instance, is based on a univariate model. Blanchard and Quah (1989) method, in turn, decompose the errors process into temporary and permanent shocks for a bivariate VAR. However, for a three variables VAR, as also considered in this study, Enders (2004) recommends that the application of this method would be complex and might generate misleading interpretations in comparison to Cholesky.

Following the formula proposed by Sims (1986) and Bernanke (1986)[87] the structural model I in (5.3.1) can be identified by imposing one restriction and the structural model II in (5.3.6) can be recovered by imposing three restrictions. For the structural model I in (5.3.1), firstly, it is assumed that the ratio of net foreign debt to GDP has contemporaneous effects on exchange rate volatility due to the increase in default risks, but exchange rate volatility does not have contemporaneous effects on that ratio so that $e_{2t} = \varepsilon_{dt}$ and $\beta_{21}(0) = 0$ in (5.3.1). Corporate financial results (balance sheet effects) are only affected after a lag period in the case of an exchange rate shock, whereas exchange rate volatility is affected contemporaneously by a debt shock. Similar reasoning can also be applied to structural model II in (5.3.6) which includes the interest rate or international reserves to form a three vector VAR. The same as in (5.3.1), once again $\beta_{21}(0) = 0$ in (5.3.6) as D_t in (5.3.6) is not affected by contemporaneous exchange rate shocks. Furthermore, the monetary authorities do not react contemporaneously to stabilise the exchange rate volatility as it takes some time to realise what is really happening in the market before intervening. If there is a time lag for a monetary intervention, then the Z_t in (5.3.6) is only affected by lagged exchange rate shocks so that $\beta_{31}(0) = 0$. Finally, D_t in (5.3.6) is not influenced by monetary intervention shocks as balance sheet effects are a long-run phenomenon and $\gamma_{23}(0) = 0$. Thus, the three restrictions assumed for structural model II in (4.6) are $\beta_{21}(0) = 0$, $\beta_{31}(0) = 0$ and $\gamma_{23}(0) = 0$.

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[87] Sims (1986) and Bernanke (1986) suggest the formula $(n^2 - n)/2$ to impose restrictions on the structural model where n is the unknown estimates in the matrix H. Note, nevertheless, that due to non-linear relationships in the system, this is considered a necessary condition, but not sufficient, for the exact identification.

Given expressions (5.3.5), (5.3.9), and based on the restrictions by a Choleski decomposition, the structural errors of structural model I and structural model II can be recovered from the representations (5.3.10) and (5.3.11), respectively[88].

The structural errors of model I are:

$$\begin{aligned} \varepsilon_{vt} &= e_{1t} + \phi_{12}(0)e_{2t} \\ \varepsilon_{dt} &= e_{2t} \end{aligned} \qquad (5.3.10)$$

as $\beta_{21}(0) = 0$.

Finally, the structural errors of model II are:

$$\begin{aligned} \varepsilon_{vt} &= e_{1t} + \phi_{12}(0)e_{2t} + \gamma_{13}(0)[e_{3t} + \phi_{32}(0)e_{2t}] \\ \varepsilon_{dt} &= e_{2t} \\ \varepsilon_{zt} &= e_{3t} + \phi_{32}(0)e_{2t} \end{aligned} \qquad (5.3.11)$$

so that,

$$\Delta = \left[1 + \phi_{12}(0)\gamma_{23}(0)\beta_{31}(0) + \beta_{21}(0)\phi_{32}(0)\gamma_{33}(0) - \beta_{31}(0)\gamma_{13}(0) - \beta_{21}(0)\phi_{12}(0) - \phi_{32}(0)\gamma_{23}(0)\right]$$

$\Delta = 1$, as $\beta_{21}(0) = 0$, $\beta_{31}(0) = 0$ and $\gamma_{23}(0) = 0$.

[88] Note that the structural errors are recovered from the following representation in matrix form: $Eee' = EH^{-1}\varepsilon\varepsilon'\left(H^{-1}\right)' = H^{-1}E\varepsilon\varepsilon'\left(H^{-1}\right)'$, such that $Eee' = \sum_e$ is the variance/covariance of the regression errors and $E\varepsilon\varepsilon' = \sum_\varepsilon$ is the variance/covariance matrix of the structural errors. Thus, $E\varepsilon\varepsilon' = HEee'H'$ or $\Sigma_\varepsilon = H\Sigma_e H'$. $Eee' = \sum_e$ is a symmetric matrix.

A vector moving average-VMA can also be derived from (5.3.3) and (5.3.8). By iterating backwards for the reduced-forms (5.3.3) and (5.3.8), the following vector moving average (VMA) representations based on regression errors emerge:

The VMA model I in matrix form is defined as:

$$\begin{bmatrix} V_t \\ D_t \end{bmatrix} = \begin{bmatrix} \bar{V}_t \\ \bar{D}_t \end{bmatrix} + \sum_{k=1}^{\infty} \begin{bmatrix} a_{11}(k) & a_{12}(k) \\ a_{21}(k) & a_{22}(k) \end{bmatrix}^k \begin{bmatrix} e_{1t-k} \\ e_{2t-k} \end{bmatrix} \qquad (5.3.12)$$

The VMA model II in matrix form is defined as:

$$\begin{bmatrix} V_t \\ D_t \\ Z_t \end{bmatrix} = \begin{bmatrix} \bar{V}_t \\ \bar{D}_t \\ \bar{Z}_t \end{bmatrix} + \sum_{k=1}^{\infty} \begin{bmatrix} b_{11}(k) & b_{12}(k) & b_{13}(k) \\ b_{21}(k) & b_{22}(k) & b_{23}(k) \\ b_{31}(k) & b_{32}(k) & b_{33}(k) \end{bmatrix}^k \begin{bmatrix} e_{1t-k} \\ e_{2t-k} \\ e_{3t-k} \end{bmatrix} \qquad (5.3.13)$$

Nonetheless, the $\{e_{it}\}$ sequences in (5.3.12) and (5.3.13), $i=(V_t, D_t, Z_t)$, are composites of structural shocks ε_{it} defined as in (5.3.4) and in (5.3.9), respectively, and regression errors expressions in matrix form can be derived.

From (5.3.4) the expression for regression errors for the model I in matrix form is as follows:

$$\begin{bmatrix} e_{1t} \\ e_{2t} \end{bmatrix} = [1/1-\beta_{21}(0)\phi_{12}(0)] \begin{bmatrix} 1 & -\phi_{12}(0) \\ -\beta_{21}(0) & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_{vt} \\ \varepsilon_{dt} \end{bmatrix} \qquad (5.3.14)$$

From (5.3.9) the expression for regression errors for the model II in matrix form is as follows:

$$\begin{bmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \end{bmatrix} = [1/\Delta] \begin{bmatrix} 1 & -\phi_{12}(0) & -\gamma_{13}(0) \\ -\beta_{21}(0) & 1 & -\gamma_{23}(0) \\ -\beta_{31}(0) & -\phi_{32}(0) & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_{vt} \\ \varepsilon_{dt} \\ \varepsilon_{zt} \end{bmatrix} \qquad (5.3.15)$$

By substituting (5.3.14) into (5.3.12) and (5.3.15) into (5.3.13), the structural vector moving average representations emerge for the model I and the model II, respectively:

$$\begin{bmatrix} V_t \\ D_t \end{bmatrix} = \begin{bmatrix} \bar{V}_t \\ \bar{D}_t \end{bmatrix} + \sum_{k=1}^{\infty} \begin{bmatrix} \theta_{11}(k) & \theta_{12}(k) \\ \theta_{21}(k) & \theta_{22}(k) \end{bmatrix}^k \begin{bmatrix} \varepsilon_{vt-k} \\ \varepsilon_{dt-k} \end{bmatrix} \qquad (5.3.16)$$

where $\theta_i = [A^k/1-\beta_{21}(0)\phi_{12}(0)] \begin{bmatrix} 1 & -\phi_{12}(0) \\ -\beta_{21}(0) & 1 \end{bmatrix}$.

$$\begin{bmatrix} V_t \\ D_t \\ Z_t \end{bmatrix} = \begin{bmatrix} \bar{V}_t \\ \bar{D}_t \\ \bar{Z}_t \end{bmatrix} + [1/\Delta] \sum_{k=1}^{\infty} \begin{bmatrix} \delta_{11}(k) & \delta_{12}(k) & \delta_{13}(k) \\ \delta_{21}(k) & \delta_{22}(k) & \delta_{23}(k) \\ \delta_{31}(k) & \delta_{32}(k) & \delta_{33}(k) \end{bmatrix}^k \begin{bmatrix} \varepsilon_{vt-k} \\ \varepsilon_{dt-k} \\ \varepsilon_{zt-k} \end{bmatrix} \qquad (5.3.17)$$

where $\delta_i = [B^k/\Delta]\begin{bmatrix} 1 & -\phi_{12}(0) & -\gamma_{13}(0) \\ -\beta_{21}(0) & 1 & -\gamma_{23}(0) \\ -\beta_{31}(0) & -\phi_{32}(0) & 1 \end{bmatrix}$.

Based on forecast error variance decompositions, the structural VMA still allows for expressions which denote impacts of a variable shock on its own variance and on the variance of another variable. By taking the square of the forecast error $[(X_{t+k} - EX_{t+k})^2 \textit{ and } (Y_{t+k} - EY_{t+k})^2]$ for each vector in (5.3.16) and in (4.17), the following expressions can be derived for both structural models:

$$\frac{\sigma_d^2\left[\theta_{12}(0)^2 + \theta_{12}(1)^2 + \ldots + \theta_{12}(5)^2\right]}{\sigma_v(6)^2} \qquad (5.3.18)$$

$$\frac{\sigma_d^2\left[\theta_{32}(0)^2 + \theta_{32}(1)^2 + \ldots + \theta_{32}(5)^2\right]}{\sigma_z(6)^2} \qquad (5.3.19)$$

$$\frac{\sigma_v^2\left[\theta_{21}(0)^2 + \theta_{21}(1)^2 + \ldots + \theta_{21}(5)^2\right]}{\sigma_d(6)^2} \qquad (5.3.20)$$

The expressions (5.3.18), (5.3.19) denote, respectively, the proportional impact of a foreign debt/GDP (d) shock on the exchange rate (v) variance and on interest rate/international reserves (z) variance for a six-quarter time horizon. The (5.3.20) expression denotes, in turn, the proportion of an exchange rate (v) shock on foreign debt/GDP (d) variance. Note, in addition, that the expressions (5.3.18), (5.3.19) and (5.3.20) enable the investigation of the three main hypotheses of this chapter: (1) the effect of increasing risks due to high levels of foreign debt on the foreign exchange market, (2) spillover effects generated by monetary authority interventions, and (3) the possibility of balance sheet changes taking place.

Finally, a further hypothesis, the impact of monetary authority interventions following exchange rate shocks can be analysed by expression (5.3.21) as follows:

$$\frac{\sigma_v^2\left[\theta_{31}(0)^2+\theta_{31}(1)^2+\ldots+\theta_{31}(3)^2\right]}{\sigma_z(4)^2} \qquad (5.3.21)$$

Basically, (5.3.21) reflects the use of monetary policy tools to control eventual instabilities in the foreign exchange market. This policy becomes evident by observing either a more volatile interest rate (*ir*) or international reserves (*res*) or both in the case of turbulence in exchange rates.

Tables 5.4 and 5.5 display the computations of variance decompositions based on (5.3.18) and on (5.3.19) from the two VARs:

Table 5.4 – Percentage accounted for by Foreign Debt Shocks - Industrialised Market Economies -

Horizon	Australia				Canada			
	Model I	Model II			Model I	Model II		
	v*	v*	ir	res	v*	v*	ir	res
0	0.35	1.12	0.00	0.00	0.08	0.078	0.00	0.00
1	9.49	11.6	0.12	0.05	0.14	0.078	0.41	0.11
2	9.99	12.1	1.95	1.82	0.31	0.37	0.99	0.14
3	13.8	15.1	1.89	1.82	1.33	1.70	5.98	2.06
4	14.5	17.1	1.91	2.01	1.62	1.75	5.98	2.18
5	15.0	17.1	1.90	1.99	1.62	1.77	5.97	2.20
6	15.7	17.7	1.89	2.05	1.82	1.94	6.46	2.45

* v : nominal exchange rate.

Table 5.5 – Percentage accounted for by Foreign Debt Shocks - Emerging Market Economies -

Horizon	Mexico				Turkey			
	Model I	Model II			Model I	Model II		
	v*	v*	ir	res	v*	v*	ir	res
0	8.62	8.71	0.00	0.00	0.014	0.019	0.00	0.00
1	20.4	19.0	0.08	0.06	1.34	3.52	2.93	8.83
2	20.6	18.7	1.83	2.98	2.00	3.37	2.92	13.4
3	20.7	18.4	5.06	18.6	2.50	3.67	6.89	15.9
4	22.0	19.5	5.98	18.1	2.92	4.43	6.82	15.8
5	22.1	20.0	6.06	19.1	2.94	4.72	6.82	15.9
6	22.1	20.1	6.11	19.2	2.96	4.71	6.8	15.9

* v : nominal exchange rate volatility.

Tables 5.6 and 5.7 demonstrate the effect of an exchange rate shock on the ratio of net foreign debt to GDP (*debt*), on the interest rate (*ir*) and on international reserves (*res*) in accordance with (5.3.20) and (5.3.21), respectively:

Table 5.6 – Percentage accounted for by ER Volatility Shocks - Industrialised Market Economies -

Horizon	Australia				Canada			
	Model I	Model II			Model I	Model II		
	debt	debt	ir	res	debt	debt	ir	res
0	0.00	0.00	0.00	0.00	0.08	0.00	0.00	0.00
1	0.01	0.03	2.81	1.26	0.01	0.002	0.02	0.32
2	1.33	2.03	2.95	1.09	0.02	0.03	0.18	2.10
3	1.88	2.37	3.14	2.91	1.58	1.47	0.35	8.03
4	2.17	2.49	2.88	2.72	2.17	2.16	0.39	8.33
5	2.48	2.83	2.91	2.80	2.24	2.20	0.40	8.23
6	2.64	3.00	2.78	2.73	2.64	2.67	0.46	8.44

Table 5.7 – Percentage accounted for by ER Volatility Shocks - Emerging Market Economies -

Horizon	Mexico				Turkey			
	Model I	Model II			Model I	Model II		
	debt	debt	ir	res	debt	debt	ir	res
0	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
1	0.35	0.51	1.01	0.04	0.16	0.08	4.00	0.38
2	0.38	0.53	1.49	2.47	2.98	4.34	3.88	1.61
3	0.34	0.49	1.30	2.37	4.77	6.17	3.92	3.51
4	0.37	0.68	1.30	2.46	5.48	7.00	3.82	3.51
5	0.38	0.68	1.33	2.50	5.48	7.03	3.84	5.37
6	0.37	0.67	1.32	2.46	5.48	7.01	3.83	5.54

For industrialised market economies, the results of the variance decomposition reveal that a foreign debt shock seems to produce significant impacts on the exchange rate volatility only in Australia (see Table 5.4). In particular, the exchange rate volatility in Australia is significantly more sensitive to a foreign debt shock in comparison to Canada. In terms of percentages, Table 5.4 shows that the exchange rate volatility in Australia responds by over 10% to a foreign debt shock after one time period, whereas in Canada this percentage is substantially lower. Moreover, Table 5.6 reveals that an exchange rate volatility shock does not seem to spill over significantly to the interest rate and international reserves. The interest rate and international reserves volatility do not grow at the same proportion after an exchange rate volatility shock supposedly attributed to a foreign debt shock. The Australian monetary authorities seem to allow the exchange rate to fluctuate more freely avoiding frequent interventions in the foreign exchange market. This conclusion meets the definition of "benign neglect" emphasised by Hausmann *et al.* (2001)[89]. The low exchange rate volatility in Canada after a foreign debt shock, in turn, may be explained by a more active monetary intervention policy by the Canadian authorities via international reserves (see Tables 5.4 and 5.6). Table 5.6 shows that the exchange rate volatility shock is usually followed by an increase in international reserves volatility in Canada. However, this result is fragile as it is not possible to establish a robust relationship between the exchange rate volatility and foreign debt shocks in Canada. Table 5.4 does not show a significant impact of foreign debt on exchange rate volatility so that the increase in international reserves volatility displayed in Table 5.6 cannot be necessarily attributed to a monetary intervention.

The feedback effect determined by an exchange rate shock on foreign debts does not appear to have significant relevance to support Bernanke, *et al.* (1999), for instance. The percentages found are very low so that a plausible hypothesis for any balance sheet change may perhaps be attributed to some changes in debt

[89] See also Calvo and Reinhart (2001) for the same result for Australia.

portfolio due to an exchange rate volatility shock[90]. It is also worth highlighting that the impact of a foreign debt shock on exchange rate volatility occurs at least after one time period. This result confirms the Choleski decomposition hypotheses assumed in the chapter.

For emerging market economies, a foreign debt shock causes a stronger impact on exchange rate volatility in Mexico in comparison to Turkey (see Table 5.5). This result may be explained by exchange rate regime differences adopted by the countries. Turkey had a fixed exchange rate regime for most of the sample period. The managed and fixed exchange rate regime adopted by both countries during part of the sample period implied necessarily the intense use of international reserves to offset any eventual exchange rate turbulence. Although the results in Table 5.7 are partly supportive of an exchange rate volatility shock spilling over significantly on the interest rate or international reserves, there is a possibility that the increase in international reserves volatility in both countries comes about from an active intervention policy of the monetary authorities in the foreign exchange market. Note that a foreign debt shock accounts for more than 15% of the international reserves volatility in Mexico and in Turkey and intense monetary interventions were required (see Table 5.5). Finally, the results also reveal some evidence of effects of an exchange rate shock on the ratio of foreign debt to GDP in Turkey (see Table 5.7). Thus, the possibility of a balance sheet change taking place via changes in debt portfolios seems to be more evident in Turkey.

In general, the results produced by the variance decomposition do reveal support for effects of foreign debt on the exchange rate volatility. However, the volatility of the ratio of debt

[90] In practice, the method used in this study does not allow us to measure currency mismatches and thus can only reach limited conclusions about balance sheet effects.

to GDP seems to have a more significant impact on exchange rate volatility in financially fragile, emerging market economies. The feedback effects denoted by balance sheet changes do not demonstrate substantial evidence.

The VMA representation as in (5.3.17) can also be used to trace out the effects of ε_{dt} and ε_{vt} shocks on the entire time path of the $\{v_t\}$, $\{d_t\}$, $\{z_t\}$ sequences. Each element $\phi_{ij}(k)$ in (5.3.17) denotes impact multipliers that inform the response of a one-unit change of ε_{it} on the $\{i_t\}$. The entire set of coefficients in (5.3.17) is called the impulse response function and can be plotted against k to derive the behaviour of the $\{v_t\}$, $\{d_t\}$, $\{z_t\}$ series in response to shocks[91].

Figures 5.2, 5.3 and Figures 5.4, 5.5 trace out the effects of a ε_{it} shock on the time path of the $\{v_t\}$, $\{d_t\}$, $\{z_t\}$ sequences for industrialised market economies and emerging market economies respectively for a forty five quarters-time horizon in the horizontal axis.

[91] Note that (5.3.16) and (5.3.17) imply orthogonalised impulse response functions.

Figure 5.2 – Responses to ε_{dt} and ε_{vt} Shocks (Australia)

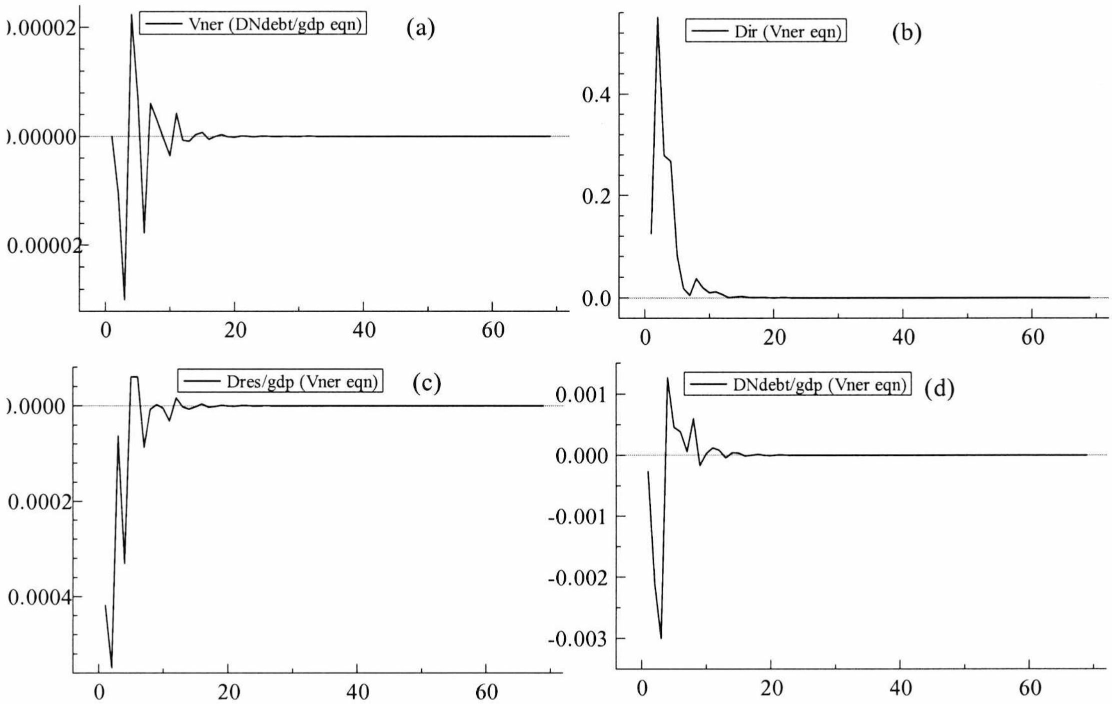


Figure 5.3 – Responses to ε_{dt} and ε_{vt} Shocks (Canada)

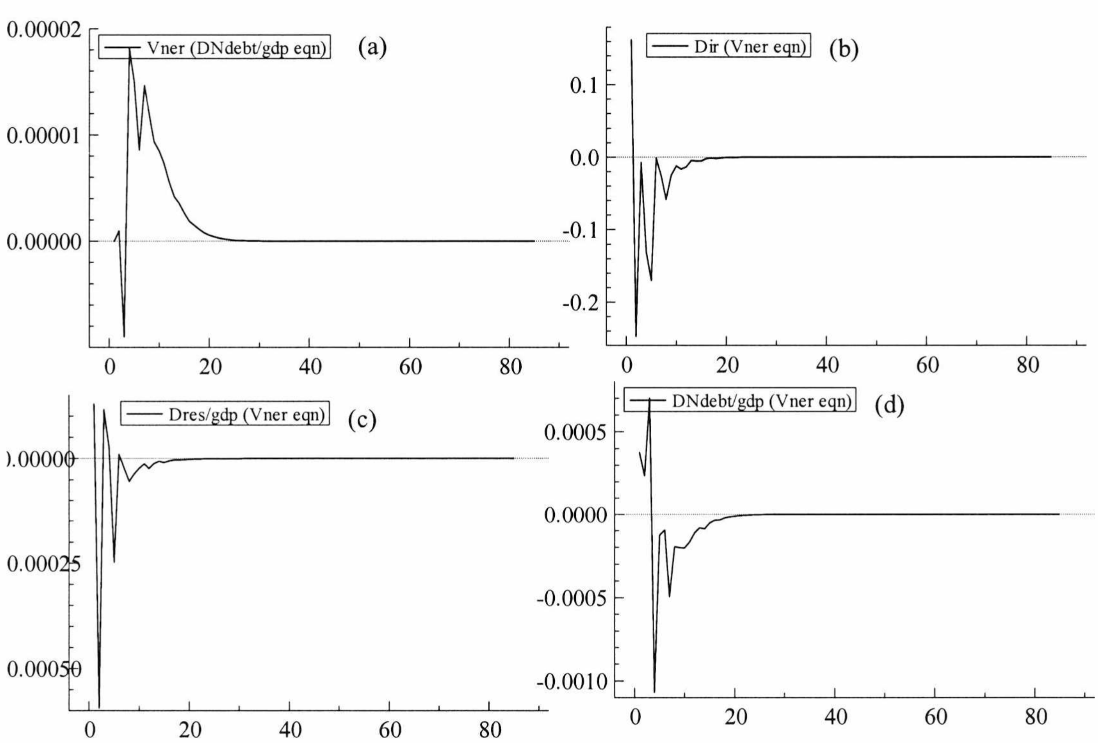


Figure 5.4 – Responses to ε_{dt} and ε_{vt} Shocks (Mexico)

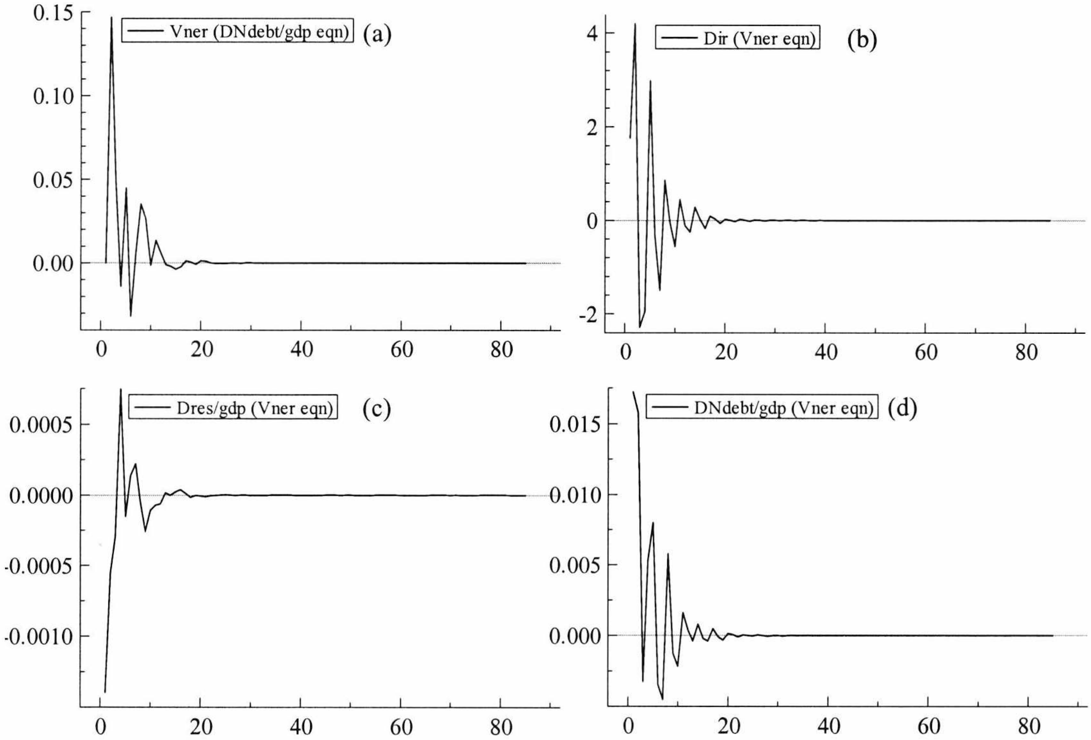
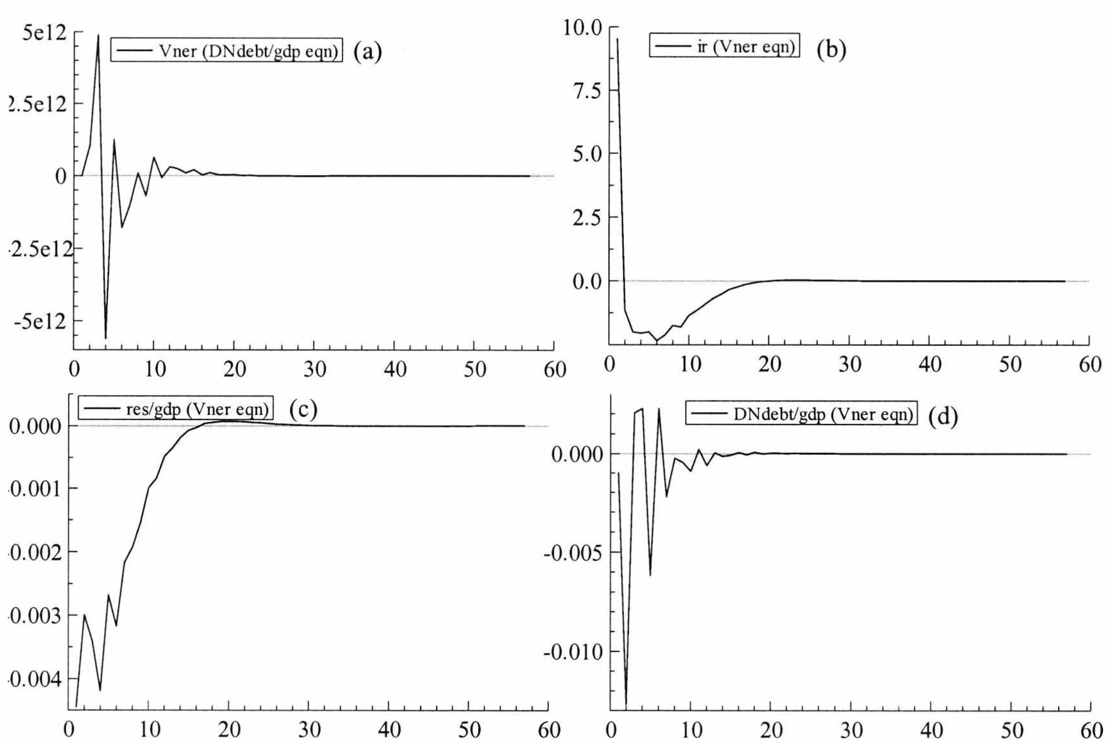


Figure 5.5 – Responses to ε_{dt} and ε_{vt} Shocks (Turkey)



The exchange rate volatility in Australia responds to a foreign debt shock after one quarter by initially showing a volatility decline (see Figure 5.2a); then the volatility in the exchange rate tends to increase after the second quarter. From the ninth quarter onwards a stabilising behaviour takes place as a result of supposedly lagged effects of interest rates (see Figure 5.2b) and some impacts from international reserves (see Figure 5.2c). In particular, the impulse response functions demonstrate evidence of a foreign debt shock generating lagged volatility in the exchange rate, but monetary interventions denoted by shocks on interest rates calm down the foreign exchange market from the ninth quarter onwards[92]. The shocks on foreign debts seem to show a similar behaviour in comparison to exchange rate volatility after a one-unit exchange rate volatility shock, but the impact on foreign debt tends to last a little shorter as the stabilisation comes after the sixth quarter. Debt portfolio adjustments may be taking place over the period (see Figure 5.2d).

The impulse response functions for Canada demonstrate that foreign debts produce impacts on exchange rate volatility after one period (see Figure 5.3a). These effects seem to last for more than nine quarters and the exchange rate volatility tends to stabilise as a result of lagged effects of international reserves. Figure 5.3c shows the impact on international reserves from the exchange rate volatility that may be associated with a monetary intervention policy. Nevertheless, Figure 5.3b reveals that the interest rate seems to be one of the monetary intervention tools as there is an increase in its level after the second quarter. The foreign debt is affected by the exchange rate volatility after the second quarter and it may

[92] Previous results based on variance decomposition for Australia revealed irrelevant monetary interventions.

demonstrate some evidence of debt portfolio changes by firms (see Figure 5.3d).

The impact of foreign debts on the exchange rate volatility in Mexico, in turn, is more lasting and significant in comparison to industrialised market economies. The exchange rate volatility is increased after the first quarter and remains unstable until the 17^{th} quarter (see Figure 5.4a). This longer time horizon may be a result of financial fragilities frequently associated with an emerging market economy along with the constraints imposed by international credit markets. Figures 5.4b and 5.4c provide evidence of lasting increases in variations for interest rates and international reserves which may be interpreted as the more intense monetary strategies adopted in Mexico to stabilise the exchange rate volatility. In particular, the movement of interest rate seems to follow a similar path demonstrated by the exchange rate volatility due to foreign debt shocks. The feedback effect on the foreign debt after an exchange rate shock also reveals more lasting behaviour in comparison to industrialised market economies and it may reveal some debt portfolio adjustment (see Figure 5.4d).

Finally, and similar to Mexico, the impulse response function for Turkey reveals a lasting effect of a foreign debt shock on exchange rate volatility (see Figure 5.5a). The increase in exchange rate volatility due to a foreign debt shock extends to the 15^{th} quarter. This behaviour does not demonstrate to be followed by increases in the interest rate or significant impacts on international reserves as it would be expected (see Figures 5.5b and 5.5c). Once again, the longer impact of a foreign debt shock on the exchange rate volatility may be a consequence of financial fragilities and credit constraints faced by Turkey. Nevertheless, the exchange rate stability in Turkey does not seem to be obtained by monetary interventions in the foreign exchange market. In fact, it is not possible to have clear conclusion on how Turkey obtains stability for its exchange rate via impulse response functions. The feedback effect denoted by an increase in foreign debt changes due to exchange rate volatility follows a longer time path in comparison to industrialised market economies (see Figure 5.5d). Perhaps, some short term debt adjustments take place over the period as result of exchange rate volatility.

The interpretation of the impulse response functions suggests that foreign debts have impacts on the exchange rate volatility and that monetary interventions in the foreign exchange market seem to be a common practice among the countries investigated (except for Turkey). In particular, the results also reveal that firms, and governments, under credit constraints tend to alter their debt portfolios in a context of higher exchange rate volatility, and thus the ratio of foreign debt to GDP tends to suffer impacts for a longer period until the exchange rate is stabilised. In contrast to emerging market economies, free access to the international credit market may allow industrialised market economies to roll over their debts easily so that an increase in exchange rate volatility does not produce an additional impact on foreign debt adjustments. Hence, the possibility of balance sheet changes taking place seems to be more significant in emerging market economies which are characterised by permanent financial fragilities and higher exchange rate volatility.

5.4 - Conclusion

This chapter has investigated the hypothesis that foreign debts may impact on exchange rate volatility. The main contribution has been to examine whether foreign indebtedness leads to a significant impact on exchange rate volatility in different economic contexts: developed and emerging economies markets. The analysis explored the stochastic properties of time series using the VAR modelling approach. It also took into account the constraints imposed by international capital markets on highly indebted economies, especially on emerging market economies. At first, this circumstance indicates that government authorities need to regulate the financial system in order to avoid financial fragilities.

The results, in general, did reveal evidence that foreign debt has robust effects on exchange rate volatility in industrialised and emerging market economies. The analysis was based essentially on variance decomposition of orthogonalised errors which demonstrated significant effects for both groups of countries. This conclusion is reinforced by a suitable impulse response analysis and did support the results obtained by Devereux and Lane (2003) though their methodology was very different. Although a foreign debt shock has a stronger impact on exchange rate volatility in emerging market economies compared to developed market economies, the findings, in general, demonstrated the power of foreign debt in affecting exchange rate volatility. Moreover, as it is suggested by Hausmann *et.al.* (2001), both types of economies make use of the traditional instruments of economic policy, such as tight monetary policies or international reserves management, or both, to stabilise exchange rate volatility. The computations did allow for some conclusion on monetary interventions, so that these policies have some role in avoiding high exchange rate volatility (spillover effects). This policy procedure is more evident in emerging market economies which operate, in general, managed exchange rate regimes.

The computations based on the impact of an exchange rate volatility shock on the ratio foreign debt to GDP did also reveal weak evidence of balance sheet changes. In particular, an exchange rate volatility shock was more likely to produce sizeable percentage impacts on the ratio of net foreign debt to GDP in Turkey. An eventual instability in the exchange rate may generate more frequent changes in debt portfolios for financially fragile and credit constrained economies. The analysis based on variance decomposition and impulse response functions did not allow for a robust conclusion.

The main policy implication drawn from this study is that government authorities have to keep foreign indebtedness at levels consistent with macroeconomic stability. It is also worthy of note that the objective of this chapter has not been to establish an empirical regularity for the effect of foreign debt volatility on exchange rate volatility across industrialised and emerging market countries The setting of an empirical regularity would need an additional research effort embodying a much larger number of countries.

CHAPTER VI

Final Conclusions

This thesis has investigated a number of issues in exchange rate economics, and the main contribution to the literature is the investigation of traditional exchange rate determinants in different economic environments. The empirical strategy was to select a sample of industrialised and emerging market economies and to examine the effects of traditional macroeconomic determinants on exchange rate behaviour. The underlying idea is that these groups of economies offer different pathologies of macroeconomic variables which have distinct impacts on exchange rate movements. The traditional monetary approach to the exchange rate was used in the analysis and in conjunction with new econometric techniques aimed at increasing the explanatory power of traditional fundamentals. An additional analysis was carried out, examining the impact of foreign debts on exchange rate volatility.

Although the objective was not to cover the whole debate about exchange rate determination, the results enabled some interesting conclusions. The monetary approach to the exchange rate was tested with panel data in which a high level of heterogeneity across countries was allowed. This model represents essentially a long-run relationship between the exchange rate and its main determinants, and monetary shocks constitute the key element in the transmission mechanism. A first finding is that panel unit root tests showed that most of these variables are integrated of order one for both types of economies, industrialised and emerging. Panel cointegration tests were also applied and the results pointed towards the possibility of no long-run relationship between the exchange rate and its monetary fundamentals for both types of economies. Even so, the rejection of the cointegration hypothesis may not mean necessarily the non-existence of long-run relationships.

A further exploration using the Pooled Mean Group estimation procedure, the results for the industrialised market economies offered some support in favour of the monetary model of the exchange rate, but with a low R^2. This estimation procedure did not find support for the monetary approach in emerging market economies. Possibly the poor performance of the monetary model for emerging market economies may be attributed to the managed exchange rate regimes usually adopted by these economies. The result that the monetary approach to the exchange rate was stronger for industrialised market economies, where free float exchange rate regimes are adopted, was somewhat unexpected. Overall the poor performance of the monetary model raised the possibility of bubbles in exchange rates.

Chapter IV took up this issue. The monetary model was used to detect the presence of speculative rational bubbles driving the exchange rate away from its economic fundamentals. This analysis was constrained to countries that operate a freely floating exchange rate regime, and thus emerging market economies were not examined. Mostly emerging market economies began to operate freely floating exchange rate regimes more recently and the time series are too short to allow for a viable examination of bubbles. The objective was to investigate a special type of bubble, usually called periodically collapsing bubbles in the literature, which have collapsing periods and expanding periods. A Markov-Switching (MS) methodology was employed, being an appropriate econometric tool to analyse stochastic processes formed by multiple regimes. Two regimes were considered: the fundamental solution regime and the bubble solution regime. The results rejected the hypothesis of a periodically collapsing rational bubble driving the exchange rate away from its fundamentals. However, linearity test statistics revealed significant evidence of exchange rate non-linearity. Non-linear behaviour invites a further investigation of the presence of market frictions such as tariff and non-tariff barriers, transaction costs, etc., in affecting exchange rate movements.

In Chapter V important results were obtained from the investigation of the impact of foreign debts on exchange rate volatility. The idea was that increasing levels of un-hedged foreign debts may affect exchange rate volatility due to default risks. Higher foreign debts may lead to an increase in default risks, capital outflows and consequent impacts on exchange rate behaviour. A sample of four industrialised and emerging market economies were selected and tested via VAR modelling. Both categories of countries, industrialised and emerging, revealed some evidence of foreign debt effects on exchange rate volatility. As expected, a stronger result was found for emerging market economies which showed bigger impacts from net foreign debts on exchange rate volatility. This finding for emerging market economies is particularly interesting; given their financial fragilities, they are vulnerable to speculative attacks and to the exchange rate having to bear the brunt of adjustment to shocks.

In addition, tests also using VAR modelling were developed to examine the possibility of monetary authority interventions to stabilise exchange rate volatility from shocks arising from foreign debt. The results showed a common characteristic for emerging market economies of using at least one of these intervention tools, either the interest rate or international reserves, to stabilise exchange rate volatility. This type of policy intervention is not so evident for industrialised market economies.

Overall, although some limited support was found, the results did not provide robust evidence of a relationship between the exchange rate and its classical macroeconomic fundamentals. The exchange rate is both a relative price and a financial asset, often subject to interference from exogenous elements which drive it temporarily away from its fundamental equilibrium. The decomposition and identification of theses extraneous components is a complex project for future research. Possibly, the use of more advanced econometric methods in conjunction with better time series data across a wider range of countries may be able to contribute with new, interesting findings.

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