Exchange Rates, fundamentals and nonlinearities:
A review and some further evidence from a century of data

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Abstract

In this paper we provide an extensive review of the monetary model of exchange rate determination which is the main theoretical framework on analyzing exchange rate behaviour over the last forty years. Furthermore, we test the flexible price monetarist variant and the sticky price Keynesian variant of the monetary model. We conduct our analysis employing a sample of fourteen advanced economies using annual data spanning the period 1880-2012. We provide strong evidence of the existence of a nonlinear relationship between exchange rates and fundamentals. Therefore, we model the time-varying nature of this relationship by allowing for Markov regime switches for the exchange rate regimes. Modelling exchange rates within this context can be motivated by the fact that the change in regime should be considered as a random event and not predictable. These results show that linearity is rejected in favour of a MS-VECM specification which forms statistically an adequate representation of the data. Two regimes are implied by the model; the one of the estimated regimes describes the monetary model whereas the other matches in most cases the constant coefficient model with wrong signs. Furthermore it is shown that depending on the nominal exchange rate regime in operation, the adjustment to the long run implied by the monetary model of the exchange rate determination came either from the exchange rate or from the monetary fundamentals. Moreover, based on a Regime Classification Measure, we showed that our chosen Markov-switching specification performed well in distinguishing between the two regimes for all cases. Finally, it is shown that fundamentals are not only significant *within* each regime but are also significant for the switches *between* the two regimes.

Keywords: Monetary model; nonlinearity; fundamentals; cointegration; Markov switching model;

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1. Introduction

A long standing puzzle in international finance is the difficulty of tying floating exchange rates to macroeconomic fundamentals such as money supplies, prices, outputs and interest rates. Economic theory states that the exchange rate is determined by such fundamental variables, but floating exchange rates between countries with similar inflation rates are in fact well-approximated as random walk. Fundamental variables do not help predict future changes in exchange rates.

Meese and Rogoff (1983) were the first to establish this result. Using data from the 1970s they examined the out-of-sample fit of alternative models of exchange rates. Their main finding was that if we use standard measures of forecasting accuracy, such as the RMSE, then the forecasting accuracy increased overall when we forecast exchange rates with a simple random walk model as compared to that of the structural models of exchange rates. During the last decade, several works have managed to provide evidence for the forecasting superiority over long horizons of structural models such as the monetary model and portfolio balance models over the simple random walk model. However, these results are not robust while these models have poor forecasting performance in the short-run. Overall, it seems that the recent results do not lead to a specific model/specification and also that while one model will do well for one exchange rate may not do for another.

An extensive subsequent literature shows that robustness of these results for the post-Bretton Woods floating period by using non-linear econometric techniques, different currencies, data periodicity and samples (Cheung *et al*. 2005).[1] Then, the difficult task to tackle is to model the exchange rates using fundamental economic variables and to obtain forward exchange rate fit both in-sample and out-of-sample to

[1]Although the initial works on testing the validity of the monetary model in all its variants have led to some positive results, the results broke down when the period was extended beyond 1978 (Backus, 1984; MacDonald, 1999).

overcome the negative result of Meese and Rogoff (1983) that exchange rates and fundamentals are separated (Frankel and Rose, 1995, p.1704).

The development of cointegration theory in the mid-1980 has given a new lease of life to exchange rate modeling. This approach provides the framework to analyze the relationship between exchange rates and fundamentals from a long-run perspective. This is a natural way to model exchange rate behaviour since it has been shown that fundamentals matter in the long-run (see Mark, 1995; Chinn and Meese, 1995; Kim and Mo, 1995; MacDonald, 1999 among the very many studies). During the last twenty years a great deal of research has been done in order to examine whether the monetary model of exchange rate is a valid long-run framework to explain exchange rate determination and movements. Indeed, although these models do not have short-run predictive performance power (since they use low frequency data, their shorter forecast is one month or one-quarter ahead) they do find evidence of long-run exchange rate predictability. This literature has been exhaustive in analysis for most of bilateral exchange rates across different exchange rate arrangements and for different time periods. What have we learned from this long-run analysis? The evidence is mixed (Cheung *et al*. 2005) and we are not allowed to form some universal results for this important relationship even in the long-run. Thus, many studies have found rather negative results for the existence of cointegration between nominal exchange rates and monetary fundamentals during the post-Bretton Woods; see for example Sarantis (1994), Berben and van Dijk (1998), Killian (1999), Berkowitz and Giovannini (2001), Faust *et al*. (2003), Engel and West (2005) and Boodoukh *et al*. 2008). Other studies notably, MacDonald and Taylor (1994), Mark (1995), Kouretas (1997), Diamandis *et al*. (1998) have found substantially support for the long-run validity of the monetary exchange rate model for several major bilateral exchange rates. Additionally, MacDonald and Taylor (1994) have also shown that the monetary model can outperform the naïve random walk model especially as we move from short-run horizons to long-run horizons. Diamandis *et al*. (2000) have further shown that considering the analysis in an $I(2)$ cointegration model we can reach to additional positive results for a number of bilateral exchange rates of the Greek drachma vis-à-vis major currencies. Recently, Dal Bianco *et al*. (2012) examine the short-run forecasting of the euro-dollar exchange rate with the development of a fundamentals-based econometric model combining weekly exchange rate data with economic variables quoted at different frequencies. The analysis provides a very good forward exchange rate in-sample fit and, more importantly satisfactory out-of-sample results.

Several explanations have been offered for this mixed evidence in finding cointegration between nominal exchange rates and monetary. A first explanation that is offered is linked with the underlying domestic and foreign demand for money functions. The instability in exchange rate modeling has been emerged as a stylized fact (see for example Faust *et al*. 2003). If these functions are unstable over time then it is difficult to find a cointegrating relation that resembles the monetary model. Berben and van Dijk (1998) and Berkowitz and Giovannini (2001) show that the positive results of studies like Mark (1995) are based on the assumption that there exists a stable cointegration relationship among the variables of interest. Indeed, it is equally important with the finding of cointegration to test for the stability of this relationship over time. Diamandis *et al*. (1998, 2000) have examined the issue of stability by applying the tests of Hansen and Johansen (1993, 1999) and they further confirm the support in favour of the monetary model. In addition, these studies have also provide an economic and statistical identification of the cointegrating results using the theoretical framework developed by Johansen and Juselius (1994) and Johansen (1995) and they were able to identify one of the cointegrating relations with the monetary model.

A second explanation for the lack of support the long-run monetary model is the relatively short sample of data usually employed which in most studies cover the post-Bretton Woods flexible exchange rates experience. Shiller and Perron (1985) and Hakkio and Rush (1991) have documented that what matters for the power of the unit root and cointegration tests typically used is not the frequency of the data but the span of the data. Coupled with the evidence for the monetary model is the equally non-favourable evidence for the long-run Purchasing Power Parity for the recent float. Given that PPP is a building block of the monetary model we can argue that failure to establish a long-run relationship between exchange rates and domestic and foreign prices possible leads to a rejection of the validity of at least the monetarist variant of the monetary model. Therefore, the low power of the standard tests is considered to be a significant factor for the substantial negative evidence.

In this paper we provide further evidence on the validity of the monetary model to the exchange rate using annual data that covers the period 1880-2013. Our analysis extends earlier works by Rapach and Wohar (2002) and Sarno *et al*. (2004) since we adopt the approach of nonlinear Markov Switching Regime modeling and we estimate the flexible-price variant of the monetary model (Frenkel, 1976) and the Keynesian sticky-price variant (Dornbusch, 1976) for Australia, Belgium, Canada, Finland, France, Italy, Portugal, Spain and the United Kingdom. We estimate a Markov Switching-Vector Error Correction model (MS-VECM) based on the evidence that at least on statistically significant cointegration vector exists. As Meese and Rogoff (1991) argue linear models can be modified by allowing non-linear formulation of coefficients. Furthermore, we examine whether the importance of exchange rates and fundamentals in restoring the long-run equilibrium level implied by the exchange rate-monetary fundamentals model varies over time and whether is affected by the nominal exchange rate regime in force.

There are several important findings stem from our analysis. Our results show that nonlinearities in the relationship between the nominal exchange rate and the macroeconomic fundamentals variables are captured fairly well by the appropriately chosen estimated MS-VECM specification for each case. Furthermore, it is shown that for all the industrialized economies during fixed exchange rate regimes it is the monetary fundamentals that adjust to restore deviations from long-run equilibrium, whereas in the cases where a less restricted exchange rate regime has been in force the exchange rate adjusts to take the system back to long-run equilibrium. These alternative adjustment schemes are also reflected by the ex-post (smoothed) transition probabilities. Finally, based on a Regime Classification Measure we show that our chosen Markov-switching specification performed well in distinguishing between the two regimes for all cases.

The rest of the paper is organized as follows. Section 2 presents the literature review. In Section 3 we provide the key elements of the flexible-price monetary model and the motivation for considering time-varying fundamentals. In section 4 the Markov switching regime methodology is presented. Section 5 reports our empirical results whereas our summary and concluding remarks are given in section 6.

2. A review of the literature

The overall poor explanatory power of structural exchange rate models (see for example Frankel and Rose, 1995) provides a natural motivation to search for a model which will take into consideration some important features of nominal exchange rates and fundamentals. There are several approaches which have been developed in the last decade to provide more sophisticated models to study the relationship between the exchange rates and fundamentals. In addition there are several propositions that have been recently advanced by economists to provide more favourable evidence for the monetary model and its out-of-sample forecasting performance.

One direction of analysis deals with the low power of standard unit root and cointegration tests. Levin and Lin (1992) offered the first response to this issue, since they recognized that the use of panel methodologies can improve substantially the power of unit root and cointegration tests. Subsequently, Groen (2000, 2005), Mark and Sul (2001), Rapach and Wohar (2004) use panel data for the post Bretton Woods era and with the application of panel cointegration tests they find strong support in favour of a stable long-run relationship between nominal exchange rates and monetary fundamentals. Moreover, these studies provide evidence that the estimated monetary model provide out-sample-forecasts that are superior to those provided by a naïve random walk model. However, Rapach and Wohar (2004) question the use of such aggregate data.[2] A second response to the low power problem is the use of long spans of data, which in most case cover more than a century. Rapach and Wohar (2002) and Sarno *et al*. (2004) are among the few studies that examined the long-run validity of the monetary model using annual data that covers the period 1880-2000 for 14 industrialized countries. They show that a stable cointegration relationship

[2]Frankel and Rose (1996), Papell (1997) and Taylor and Sarno (1998) are among the studies which have used panel cointegration techniques that led to strong support in favour of long-run PPP during the recent float.

between nominal exchange rates and monetary fundamentals could be established for Belgium, Finland, France, Italy, the Netherlands, Portugal, Spain and Switzerland.[3]

A second direction to exchange rate modeling calls for the introduction of nonlinearities in the relationship between nominal exchange rates and fundamentals. Thus, Hsieh (1989), Meese (1990) and Balke and Forby (1997), Taylor and Peel (2000), Taylor *et al*. (2001) and So (2001) have documented the existence of various nonlinearities in deviations of the spot nominal exchange rate from economic fundamentals. De Grauwe (2000) further underlines the significance of changing beliefs of the economic agents as a possible source for the existence of nonlinearities. Threshold cointegration and Markov switching regimes models are the two workhorses of this strand in the literature of estimating relationships with nonlinear features.

Threshold models assume nonlinear mean reversion of exchange rates and smooth threshold dynamics. Within this framework Kilian and Taylor (2003) provide support in favour of the monetary model while their specification also exhibits superior out-of-sample performance compared to the random walk model.

Markov switching models focus on the idea of changing regimes and time-varying coefficients. Engel and Hamilton's (1990) important contribution provided evidence that a Markov-switching model of exchange rate outperforms the naïve random walk model. The intuition behind these models relies on the evidence offered by some studies that the monetary model performs well for some sub-period of the total sample but not for others (Meese, 1990) and also that there have been observed sudden regimes changes. Frydman and Goldberg (2001) show that such regime changes occur in the case of the dollar-mark exchange over the recent float. Mahavan

[3] Again Abuaf and Jorion (1990), Glen (1992), Lothian and Taylor (1996, 2000) and Taylor (2002) use long spans of data and they provide support in favour long-run PPP.

and Wagner (1999), Marsh (2000), Bessec (2003), Clarida *et al*. (2003), Taylor and Peel (2000), Taylor *et al*. (2001), Sarno *et al*. (2004) and De Grauwe and Vansteenkiste (2007) are among several studies which analyze the monetary model in a Markov-switching model for a set of main bilateral exchange rates and they provide support in favour of a fundamental model. Bacchetta and van Wincoop (2013) argued that large and frequent variations in the relationship between the exchange rate and macroeconomic fundamentals become evident when structural parameters in the economy are unknown and subject to changes. Furthermore, Frommel *et al*. (2005a,b) examine the RID variant of the monetary model within the Markov-switching approach and they show that in a two-regime model, the one regimes accurately describes the RID monetary model and additional gives significant out-of-sample forecasting performance. Ducker and Neely (2007) provided strong evidence that the Markov-switching regime models created ex ante trading rules in the foreign exchange market and delivered strong out-of-sample portfolio returns for several major currencies. Recently, Syllignakis and Kouretas (2011) using data for the CEE economies employed a Markov-switching vector error correction model which allowed for regime shifts in the entire set of parameters and the variance-covariance matrix. The main finding of the analysis was that depending on the nominal exchange rate regime in operation, the adjustment to the long run implied by the monetary model of the exchange rate determination came either from the exchange rate or from the monetary fundamentals. Moreover, based on a Regime Classification Measure, it is shown that the chosen Markov-switching specification performed well in distinguishing between the two regimes for all cases.

A third approach to deal with the poor forecasting performance of the monetary model is relied on the development of models that use very high-frequency data based on microeconomic variables linked to the structure of the market (Lyons, 2001). One of the arguments put forward in this line of research is that the failure of macroeconomic fundamentals in explaining and forecasting exchange rates can be resolved through the analysis of the microstructure of FX markets. This type of analysis focuses in the identification of certain elements of this market will help us to understand the exchange rate. The relevant studies have concluded that order flows are a significant variable in understanding and forecasting the exchange rate (see Goodhart, 1988; Lyons, 1995; Gehrig and Menkoff, 2004; Evans and Lyons 2002, 2005; Osler (2006).

The final approach to exchange rate modeling relies on the use of real-time macroeconomic data in the estimations and forecasts to evaluate the usefulness of monetary model in predicting with the same information which market participants have at each moment. Recently, Sarno *et al*. (2009) using real-time data on a broad set of economic fundamentals for five major US dollar exchange rates over the recent float they re-examine the predictive ability of exchange rate models within this framework. Their analysis leads to two key findings. First, they argue that the stylized fact of poor forecasting performance of exchange rate models may be the outcome of poor performance of model-selection criteria, rather than the lack of information content in the fundamentals. Second, they argue that the difficulty of selecting the best predictive model is largely due to frequent shifts in the set of fundamentals driving exchange rates, which can be interpreted as reflecting swings in market expectations over time. Furthermore, it is argued that the strength of the link between exchange rates and fundamentals differs across countries.

Furthermore, as Frankel (1996) argues exchange rates are separated from fundamentals because of swings in expectations about future values of the exchange rate. He provides substantial evidence that support this argument. Therefore, this approach supports the argument that exchange rate models exhibit poor performance not only because the information content of the fundamentals is deficient, but because volatile expectations and departures from rationality are likely to account for the failure of exchange rate models. Bachetta and van Wincoop (2004) developed an exchange rate model which incorporates the fact that practitioners in the foreign exchange market regularly change the weight they attach to different macroeconomic variables (see Cheung and Chinn, 2001), This fact is supported by the results of various survey studies based on the framework of a stylized rational-expectations model of exchange rate determination. The intuition in this model is that each time that an explanation is sought by rational agents for the observed exchange rate change, a macroeconomic variable is chosen for this purpose (more weight is put on it) whereas the rest of potentially significant macroeconomic variables are left out. Therefore it is argued that different observed variables may be taken as the scapegoat, so that weights attributed to economic variables change.

3. The monetary model and nonlinear characteristics of fundamentals

The monetary model of exchange rate determination is an extension of the quantity theory of money to the case of an open economy. It assumes that: (i) real income and money supply are determined exogenously; (ii) capital and goods are perfectly mobile; (iii) foreign and domestic assets are perfect substitutes; (iv) goods' prices are perfectly flexible; and (v) domestic (foreign) money is demanded only by domestic (foreign) residents. The early, flexible-price monetary model (Frenkel, 1976) relies on the twin assumptions of continuous purchasing power parity (PPP) and the existence of stable money demand functions for the domestic and foreign economies. Recent experience with flexible exchange rates has shown, however, that real exchange rates have fluctuated substantially over the years causing shifts in international competitiveness. Stickiness in prices (Dornbusch, 1976) in conjunction with the uncovered interest parity (UIP) condition are usually invoked in order to allow for short-term deviations of both the nominal and the real exchange rates from their long-run levels as determined by the PPP. Moreover, the UIP condition is necessary for the derivation of the forward-looking version of the monetary model, under which the exchange rate depends on all the expected realizations of the forcing variables, that is, the monetary aggregates and the output variables. These two approaches are considered to be subcases of the the real interest rate differential (Frankel, 1979).

Under these assumptions a typical monetary reduced form equation is obtained (see Taylor, 1995; Sarno and Taylor, 2002 and MacDonald, 2007):

$$e_t = \beta_0 + \beta_1(m_t - m_t^*) - \beta_2(y_t - y_t^*) + (i_t - i_t^*) + u_t \qquad (1)$$

where e_t is the spot exchange rate (home currency price of foreign currency); m_t denotes the domestic money supply; y_t denotes domestic income; i_t denotes domestic interest rate; corresponding foreign magnitudes are denoted by an asterisk; u_t is a disturbance error; and all variables apart from the interest rate terms, are expressed in natural logarithms.

The expected signs of the coefficients in (1) are: $\beta_1 > 0$, $\beta_2 < 0$. Different signs of the interest rate coefficients in equation (1) will be produced under imperfect substitutability between the assets of the two countries. Associated with equation (1) is a set of coefficients restrictions that are regularly imposed and tested. The most important restriction is whether proportionality exists between the exchange rate and relative monies ($\beta_1 = -\beta_2 = 1$).

As we already mentioned several papers consider parameter instability as an explanation for the poor forecasting performance of the monetary model. This instability can be explained either by policy regime changes or instabilities in the money demand (an explanation offered in the early studies of monetary model) or PPP equations or agents' heterogeneous beliefs (see Rossi, 2005, 2006). In addition, another source of the failure of monetary model to provide accurate exchange rate forecasts may be due to changes in way expectations are formed when a switch from fixed exchange rates to flexible exchange rates occur (Flood and Rose, 1995).[4]

Equation (1) implies that, if the departure from the exchange rate-monetary fundamentals relationship u_t is stationary given e_t, $m_t - m_t^*$, $y_t - y_t^* \sim I(1)$, the nominal exchange rate and the fundamentals exhibit a common stochastic trend and are cointegrated with cointegrating vector [1, -1], i.e. the proportionality hypothesis holds. Then given the Granger Representation Theorem (Engle and Granger, 1987), the nominal exchange rate and the fundamentals must possess a VECM representation in which u_t plays the part of the error correction term. We follow Sarno *et al*. (2004) and we use exactly a linear VECM representation in order to examine the relative importance of the nominal exchange rate and the fundamentals in restoring equilibrium in the long-run relationship linking exchange rate and fundamentals across different exchange rate regimes since the late 19th century. Therefore, we employ a generalization of a standard linear VECM which is capable of allowing all of the VECM parameters to change over time and to identify the various regimes that characterize the long sample periods that we examine in the present paper.

[4] Uncovered Interest parity (UIP) is frequently invoked to provide the rational expectation version of the monetary model. However, empirical evidence in favour of UIP is rather weak. Given that we use a more than a century long data we prefer to examine the validity of the monetary model in its reduced form formation given by eq. (1).

4. Econometric methodology

4.1. Johansen Multivariate cointegration technique

The estimation of the proposed MS-VECM is conducted using a two-stage maximum likelihood procedure. The first stage refers to the cointegration analysis which is based on the multivariate cointegration technique developed by Johansen (1988, 1991) and extended by Johansen and Juselius (1990) which is a Full Information Maximum Likelihood (FIML) estimation method. It makes use of the information incorporated in the dynamic structure of the model and it also estimates the entire space of the long-run relationships among a set of variables, without imposing a normalization on the dependent variable *a priori*. Although the Johansen procedure is well known we discuss it briefly in light of some recent extensions of the methodology that are applied in this paper.

Consider a ***p***-dimensional vector time series z_t with an autoregressive representation (AR) which in its error correction form is given by

$$\Delta z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \Pi z_{t-1} + \gamma D_t + \mu_0 + \mu_1 t + \varepsilon_t, \qquad t = 1, \ldots, T \tag{2}$$

where $z_t = [e, m, m^*, y, y^*, i, i_t^*]_t$ as defined in section 3, $z_{k+1}, \ldots, z_0$ are fixed and $\varepsilon_t \sim Niid_p(0, \Sigma)$. The adjustment of the variables to the values implied by the steady state relationship is not immediate due to a number of reasons like imperfect information or costly arbitrage. Therefore, the correct specification of the dynamic structure of the model, as expressed by the parameters $(\Gamma_1, \ldots, \Gamma_{k-1}, \gamma)$, is important in order that the equilibrium be revealed. The matrix $\Pi = \alpha\beta'$ defines the cointegrating relationships, β, and the rate of adjustment, α, of the endogenous variables to their steady state values. D_t is a vector of nonstochastic variables, such as centered seasonal dummies which sum to zero over a full year by construction and are necessary to account for short-run effects which could otherwise violate the Gaussian assumption, and/or intervention dummies; μ is a drift and T is the sample size.

If we allow the parameters of the model $\theta = (\Gamma_1,\ldots\ldots,\Gamma_{k-1,}\Pi,\gamma,\mu,\Sigma)$ to vary unrestrictedly, then model (2) corresponds to the *I*(0) model. The *I*(1) and *I*(2) models are obtained if certain restrictions are satisfied. Thus, the higher-order models are nested within the more general *I*(0).

It has been shown (Johansen, 1991) that if $z_t \sim I(1)$, then that matrix Π has reduced rank $r < p$, and there exist $p\mathrm{x}r$ matrices α and β such that $\Pi = \alpha\beta'$. Furthermore, $\Psi = \alpha_{\perp}^{'}(\Gamma)\beta_{\perp}$ has full rank, where $\Gamma = I - \sum_{i=1}^{K}\Gamma_i$ and $\alpha_{\perp}$ and $\beta_{\perp}$ are $p\mathrm{x}(p-r)$ matrices orthogonal to α and β, respectively.

Following this parameterization, there are r linearly-independent stationary relations given by the cointegrating vectors β and $p-r$ linearly-independent non-stationary relations. These last relations define the common stochastic trends of the system and the contribution to the various variables. By contrast the AR representation of model (2) is useful for the analysis of the long-run relations of the data.

4.2. The Markov Switching Vector Error Correction Model

The second stage amounts to the study of the dynamics of the regime switching and the stochastic processes evolved in a set of the nominal exchange rate and fundamental variables that mentioned before, we adopt the Markov Switching Vector Error Correction Model (MS-VECM), introduced in Krolzig (1997), which is a multivariate generalisation of the univariate Hamilton (1989, 1994) model. This model allows, in a multivariate context, for shifts in the stochastic volatility regime driving the foreign exchange markets. Thus, the change in regime should be considered as a random event and not predictable. In addition, the effect of these shifts must be considered when we investigate the stochastic properties of the foreign exchange market volatility and the possible links between the exchange rate and monies supplies, outputs. The usefulness of a time-varying coefficients approach against structural models with constant coefficients has been illustrated in several studies (see for example, Wolff, 1987; Schinasi and Swamy, 1989).[5] Later studies, separate the switches in mean and variance either by using two distinctive state variables for mean and variance each (Dewachter, 1997) or by using a Markov switching model with four states differing in mean or variance (Dewachter, 2001).

Consider that Δy_t is a $Tx1$ vector containing the observations for the single stationary time series $\{\Delta y_t\}$, and let $\Delta Y_t = (\Delta y_{1t}, \ldots, \Delta y_{Kt})'$, $t = 1, \ldots, T$ be the K-dimensional vector, where T is the sample size. A p-th order MS-VECM [MS-VECM(p)] model can be written as

$$\Delta Y_t = A_0(s_t) + A_1(s_t)\Delta Y_{t-1} + \ldots + A_p(s_t)\Delta Y_{t-p} + B(s_t)ect_{t-1} + u_t,\ u_t \sim NID(\mathbf{0}, \Sigma(s_t)) \quad (3)$$

where s_t is the unobservable regime, $A_0(s_t), \ldots, A_p(s_t)$ are regime-dependent autoregressive parameter matrices, $B(s_t)$ is the regime-dependent parameter matrix of the error correction term (ect), and u_t is the innovation process with a regime-dependent variance-covariance matrix $\Sigma(s_t)$. It is assumed that s_t follows an irreducible ergodic m-regime Markov process with the transition matrix

[5] See also, Engel (1994), Kim (1994), Hamilton and Susmel (1994), Hamilton and Lin (1996), Engel and Kim (2001), Lee and Chen (2006), Kanas and Kouretas (2007)

$$P = \begin{bmatrix} p_{11} & p_{12} & \cdots & p_{1M} \\ p_{21} & p_{22} & \cdots & p_{2M} \\ \cdots & \cdots & \cdots & \cdots \\ p_{M1} & p_{M2} & \cdots & p_{MM} \end{bmatrix} \tag{4}$$

The transition probabilities p_{ij} in **P** are constant, and given by

$$p_{ij} = \Pr(s_{t+1} = j \mid s_t = i), \sum_{j=1}^{m} p_{ij} = 1, \ \forall i, j \in \{1,...,m\} \tag{5}$$

Maximum likelihood estimation of the model is based on the Expectation Maximisation (EM) algorithm.[6] One can also calculate the unconditional probability that the system of the two currency is in regime *i, i* = 1,…*m*, at any given date, Pr(s_t = *i*). Also, the 'smoothed' probabilities can be obtained, representing the ex-post inference about the system being in regime *i* at date *t*. Further, one could date the regime switches. For instance, for 2 regimes, an observation is assigned to the first regime if $Pr(s_t = 1 \mid \Delta Y_T) > 0.5$, and to the second regime if $Pr(s_t = 1 \mid \Delta Y_T) < 0.5$.

5. Data and empirical results

The data consist of annual observations for the nominal exchange rate (units of foreign currency per US dollar), the money supply, real GDP and short-term interest rates for fourteen advanced economies, Australia, Belgium, Canada, Denmark, Finland, France, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, Switzerland and the United Kingdom. The corresponding money supply and real GDP for the US are denoted with an asterisk. The data spans from 1880 to 2012 except for Belgium, Finland, France, Italy, the Netherlands, Portugal and Spain for which the sample runs until 1998 which marks the end of their national currency in

[6] The EM algorithm was first developed by Dempster *et al*. (1977) and was extended by Hamilton (1989) and Krolzig (1997).

light on the formation of the Eurozone in January 1, 1999 and thus covers a number of alternative international monetary arrangements, such the gold standard, the Bretton Woods period and the current flexible exchange rate regime. Due to specific data availability in particular for the short term interest rates the exact dates for each case are as follows: Australia (1900-2012); Belgium (1880-1998); Canada (1900-2012); Denmark (1910-2012); Finland (1910-1998); France (1880-1998); Italy (1900-1998); Netherlands (1880-1998); Norway (1900-2012); Portugal (1910-1998); Spain (1910-1998); Sweden (1880-2012); Switzerland (1880-2012); United Kingdom (1880-2012) and United States (1880-2012). The nominal exchange rates, the money supplies and real GDP are obtained from the data set employed by Rapach and Wohar (2002) and the short term interest rates were obtained from Homer and Sylla (2005). All series were updated using data taken from the *International Financial Statistics* of the International Monetary Fund.[7] All variables are measured in natural logarithms.

The standard practice is to subject variables to a battery of unit root tests. There are many different tests for a unit root in the autoregressive (AR) polynomial of a univariate process that have been proposed, but the most common is the augmented Dickey-Fuller (ADF) test proposed by Said and Dickey (1983). It is based on the AR approximation of a general ARIMA process and is given in (1).[8]

$$\Delta y_t = \alpha + (\rho - 1) y_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta y_{t-j} + e_t \tag{6}$$

[7] In turn, the nominal exchange rates series are from Taylor (2002), and the money supply and real GDP series are from Bordo and Jonung (1998), Bordo *et a*l. (1998) and Bordo's Financial Crises Database https://sites.google.com/site/michaelbordo/home4.

[8] An ARIMA or autoregressive integrated moving average process assumes that a time series can be modeled in the time domain as a function of lagged values of itself and current and lagged values of the innovation or error to the process. An ARIMA(p,d,q) takes the general form $\phi(L)\Delta^d y_t + \mu = \theta(L)\varepsilon_t$ where the autoregressive lag polynomial $\phi(L) = 1 + \phi_1 L + \phi_2 L^2 + ... + \phi_p L^p$ is of order p, the order of integration is given by the differencing parameter d and the moving average polynomial $\theta(L) = 1 - \theta_1 L - \theta_2 L^2 - ... - \theta_q L^q$ if of order q.

The null hypothesis of a unit root can be tested by estimating (1) using OLS and then using a t-type test statistic to test the hypothesis $(1 - \rho) = 0$. The choice of the lag truncation parameter k is important for the small sample properties of the test because when the number of lags is greater than the true number of lags there is a decrease in the power of the test, while two few lags leads to under sized tests. There are some potential problems with unit root testing using (6), however.

The first problem is low power of the test relative to local alternatives. Elliot *et al*. (1996) (ERS) proposed an estimator that increases the power of the unit root test substantially by using a GLS detrending procedure. One can motivate the unit root tests using the DGP in (7)

$$y_t = d_t + u_t, \qquad u_t = \rho u_{t-1} + v_t \tag{7}$$

where $v_t = \varphi(L)e_t = \sum_{j=0}^{\infty} \varphi_j e_{t-j}, d_t = \zeta' z_t = \sum_{i=0}^{p} \zeta_i t^i$ for $p = 0, 1$. When estimating equation (6) the parameters of the deterministic components are estimated via OLS and are treated as nuisance parameters in the distribution of the unit root tests. By estimating these nuisance parameters using OLS the power of the test statistics is diminished. ERS propose a weighted least squares or GLS method to estimate these parameters and then detrend the data prior to testing for a unit root. For series $\{x_t\}_{t=0}^{T}$ define $(x_0^{\overline{\alpha}}, x_t^{\overline{\alpha}}) = (x_0, (1 - \overline{\alpha}L)x_t)$ for some value $\overline{\alpha} = 1 + \overline{c}/T$. The GLS detrended series is then defined as $\widetilde{y}_t \equiv y_t - \widehat{\zeta}' z_t$ where $\widehat{\zeta}$ minimizes $S(\overline{\alpha}, \zeta) = (y^{\overline{\alpha}} - \zeta' z_t^{\overline{\alpha}})'(y^{\overline{\alpha}} - \zeta' z_t^{\overline{\alpha}})$. ERS suggest imposing $\overline{c} = -7.0$ for $p = 0$ and $\overline{c} = -13.5$ for $p = 1$.[9] Testing for a unit root can then be done by estimating equation (8) using OLS and calculating a t-type test statistic as in (1), which is referred to as the DF-GLS$^{\mu}$ statistic when $p = 0$ and DF-GLS$^{\tau}$ when $p = 1$.

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[9] Cook (2006) finds that the power of the tests in finite samples under alternative DGPs can be increased with alternative values for $\overline{c}$.

$$\Delta\widetilde{y}_t = (\rho - 1)\widetilde{y}_{t-1} + \sum_{j=1}^{k}\gamma_j \widetilde{y}_{t-j} + e_{tk} \qquad (8)$$

Although low power is always a problem for unit root tests, another concern is that size distortions in the tests may be a problem because of the properties of the underlying data generating process (DGP). One source of size distortion is the presence of large and negative moving average (MA) parameters in the DGP. Schwert (1987) was one of the first to point out that standard unit root tests like the ADF are severely oversized when there are large negative MA terms in the DGP. He suggests increasing the value of *k*, the lag truncation parameter in (6) and (8), to more accurately allow the AR process in (10) to approximate the MA components in the ARIMA. We estimate ARIMA models for each of the series of interest in this study in order to gauge how serious this source of size distortion may be in our application. Table 1 displays estimation results for our series.[10]

Two features of many economic time series tend to affect the size and power of usual unit root tests. In particular, a large negative moving average root may induce size distortions, while a large autoregressive root may result in low power. When this is the case it is preferred to apply the MZ_α, MZ_t, MSB and the MPT tests due to Ng and Perron (2001), which are precisely designed to overcome both size distortion and low power problems when the data are characterized by these features. These tests are extensions of the *M* tests of Perron and Ng (1996) that use Generalized Least Squares (GLS) detrending of the data, together with a modified information criterion for the selection of the truncation lag parameter.

Ng and Perron (2001) have developed a modified information criterion that chooses *k* in (6) or (8) in a way that mitigates the size distortion in unit root tests. It is

[10] We used the Box-Jenkins procedure to identify several candidate models for each series and then chose the best fitting model based on residual serial correlation tests, significance of the parameter estimates and R^2.

based upon an autoregressive estimate of the long-run variance of y_t, denoted s^2_{AR}. This estimate is calculated as

$$s^2_{AR} = \frac{\widehat{\sigma}^2_k}{[1-\gamma(1)]^2} \tag{9}$$

where $\gamma(1) = \sum_{i=1}^{k} \gamma_i$ and $\widehat{\sigma}^2_k = (T-k)^{-1}\sum_{t=k+1}^{T} \widehat{e}^2_{tk}$ and γ_i and $\{\widehat{e}_{tk}\}$. The parameters can all be estimated from equation (8) using OLS.[11] The modified information criteria (MIC) is given as

$$MIC(k) = \ln(\widehat{\sigma}^2_k) + \frac{C_T(\tau_T(k)+k)}{T-k_{\max}} \tag{10}$$

where $\tau_T(k) = (\widehat{\sigma}^2_k)^{-1}\widehat{\rho}\sum_{t=k_{\max}+1}^{T} \widetilde{y}^2_{t-1}$ and $k_{\max}$ is the largest lag truncation considered. When $C_T = \ln(T-k_{\max})$, equation (10) represents the modified Bayesian information criterion (MBIC) and when $C_T = 2$ it is the modified Akaike information criterion (MAIC).

Ng and Perron (2001) also suggest using three tests that have less size distortion in the presence of MA errors than standard tests. These tests are MZ_ρ, MZ_t, and MSB, collectively referred to as the M-tests. The tests are calculated from estimates of (8) as follows:

$$MZ_\rho = (T^{-2}\widetilde{y}^2_t - s^2_{AR})(2T^{-2}\sum_{t=1}^{T}\widetilde{y}^2_{t-1})^{-1} \tag{11}$$

and

$$MSB = \left[\frac{T^{-2}\sum_{t=1}^{T}\widetilde{y}^2_{t-1}}{s^2_{AR}}\right]^{\frac{1}{2}} \tag{12}$$

[11] Perron and Qu (2007) suggest that small sample power can be improved if the parameters used to construct the estimate of the long-run variance are estimated from equation (6) rather than (8).

and $MZ_t = MZ_\rho \times MSB$.

Finally, since it has been shown that the standard DF and PP unit root tests are biased towards the acceptance of the unit root hypothesis we also apply the Kwiatkowski *et al*. (1992) KPSS test for the null hypothesis of level or trend stationarity against the alternative of non-stationarity and these additional results will provide robust inference. The KPSS test has two components: in the first the null hypothesis is the stationarity of series in level; in the second the null hypothesis is that of trend stationarity. We implement the second test only when the first hypothesis is rejected. The results of the unit root and stationarity tests are presented in Table 1. The results unambiguously lead to the conclusion that we are unable to reject the null hypothesis of non-stationarity based on the DF-GLS$_u$ and MZ_a^{GLS} and MZ_t^{GLS} tests and we reject the null hypothesis of stationarity with the KPSS test for the levels of all series. However, when we take the first difference of each variable then all tests indicate that these series are I(1) processes.

Given the low power of the unit root tests against alternative hypotheses it is important that we also test for structural breaks in the time series when analysing the stochastic properties of the nominal exchange rates, the relative money supplies and the relative real output. This seems appropriate given the variety of exchange rate regimes that our data covers. In that respect, we employ the Zivot and Andrews (1992) test with one structural break. The endogenous structural break test of Zivot and Andrews (1992) is a sequential test which utilizes the full sample and uses a different dummy variable for each possible break date. This test has several desirable properties: (a) it determines the structural breaks "endogenously" from the data, (b) its null distribution is invariant to level shifts in a variable, and (c) it is easy to interpret;

by including breaks under both the null and alternative hypotheses, a rejection of the null hypothesis of a unit root implies unambiguously trend stationarity.

Furthermore, for reasons of comparison and robustness we also apply the recently developed by Perron and Rodriguez (2003) detrended GLS unit root tests against the alternative of stationarity around a structural break, which is an extension of the Elliott *et al*. (1996) detrended GLS unit root tests we used above. As in the Zivot and Andrews (1992) test the structural change is allowed to occur at an unknown point of time. The results are also shown in Table 2 and in all cases there is no evidence of one or two structural breaks in nominal exchange rate, relative money supplies and relative real outputs.

We conclude our unit root testing by applying the unit root tests developed by Kapetanios *et al.* (2003). These tests are constructed within the nonlinear STAR framework and they have better properties compared the Dickey-Fuller test. Specifically, Kapetanios *et al*. (2003) analyze the implications of the existence of a specific type of nonlinear dynamics for unit root testing procedures. They develop a test for the null hypothesis of a unit root process against an alternative of a nonlinear exponential smooth transition autoregressive (ESTAR) process which is globally stationary. Furthermore, this testing procedure has been designed to have power against this alternative ESTAR process. The results of the application of this test are also reported in Table 2. We consider the case of a constant and a constant and a linear trend in the series. For each of our series we are unable to reject the null of a unit root in favour of nonlinearity at conventional levels of significance.

Based on the evidence from the unit root and stationarity tests as well as form the structural break tests we conclude that for each case $e_t \sim I(1), (m_t - m_t^*) \sim I(1), (y_t - y_t^*) \sim \mathrm{I}(1)$ and $(i_t - i_t^*) \sim I(1)$ and therefore we can tests the full version of the monetary model by applying the Johansen (1988, 1991) and Johansen and Juselius (1990) multivariate cointegration technique described in Section 3. Table 3 reports the cointegration results based on the trace test proposed by Johansen (1988, 1991). Our overall findings show that we are able to identify one stable and statistically significant cointegrating vector for each bilateral nominal exchange rate.[12]

To address the issue of volatility regime switching and to discriminate between low and high volatility regime in the relationship between the nominal exchange rate and fundamentals, we estimate and test for an MS-VECM given by (3). In principle given that during the period under examination several exchange rate regimes have been adopted we could consider three potential regimes; regime 1 which covers flexible exchange rates (high volatility regime); regime 2 which covers managed float or peg exchange rates (the medium volatility regime) and regime 3 which covers fixed exchange rates (low volatility regime). Table 5 reports the estimated coefficients of the proposed MS-VECM along with the necessary test statistics for evaluation of the adequacy of the estimated model.[13] The Likelihood Ratio test for the null hypothesis of linearity is statistically significant and this suggests that linearity is strongly rejected. This is a nonstandard LR test due to Davies (1987). This outcome is reinforced from the AIC and HQIC criteria.

The estimation of the MS-VECM specification was conducted with the adoption of the "bottom-up" procedure (Krozlig, 1997), which was designed to determine the appropriate MS(m)-VECM(p) model for each CEE country. Table 5

[12] The estimated cointegration coefficients have the correct sign and reasonable magnitude as predicted by the monetary model. Furthermore, when testing for the proportionality hypothesis this is found to hold for Australia, Belgium, France, Italy, Netherlands, Spain and Switzerland. To save space these results are available upon request.

[13] To save space we do not report all of our MS-VECM. We only report the estimated equilibrium correction coefficients for each equation of the MS-VECMs.

presents our results for the choice of the appropriate MS-VECM specification. Specifically for all cases we estimated a MSIAH(2)-VECM(p) specification. The Likelihood Ratio test for the null hypothesis of linearity (LR1) was statistically significant for all cases which suggest that linearity is strongly rejected. This is a nonstandard LR test advanced by Davies (1987).[14] This outcome was reinforced by the AIC, SIC and HQIC criteria. Furthermore, Table 5 reports the results from two Likelihood Ratio tests (LR2 and LR3) which were used to choose the appropriate model specification. Based on these two Likelihood Ratio test statistics (Krolzig, 1997, p. 135-136) the most appropriate model within this class of MS-VECM model was the MSIAH(2)-VECM(p).[15] Additional evidence for the appropriateness of the estimated model was given by the standardized residuals which reveal no evidence of serial correlation, heteroskedasticity or substantial departures from normality.[16] Based on these estimates we argue that they are in favour of a non-linear relationship, between exchange rates and macroeconomic fundamentals.[17]

In Table 3 we report the regime-dependent equilibrium correction coefficients. For Australia, in Regime 1 we observe that the coefficients of the relative money and

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[14] The results from the LR test from Davies (1987) are reported with caution. It is argued that since the Markov regime switching model has both a problem of nuisance parameters and a problem of 'zero score', under the null hypothesis, we cannot use the χ^2 distribution to determine the significance of the LR test (Garcia, 1988). Therefore, Ang and Bekaert (2002a,b) have suggested alternative LR tests for the case in which the regularity conditions of the Davies (1987) test are not met. However, given the support obtained by the AIC, SIC and HQIC information criteria, we argue that the rejection of the linear model in favour of a MS specification is robust.

[15]The number of regimes is 2 since the estimation of models with 3 regimes is not feasible due to large number of parameters. The MSIH(2)–VECM(p) specification is given by the following expression: $\Delta X_t = v(z_t) + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \Pi(z_t) X_{t-1} + u_t$ where $u_t \left| z_t \right. \sim NID(0, \Sigma[z_t])$ and $z_t \in \{1,2\}$.

[16] Following Sarno *et al*. (2004), we further evaluated the goodness-of-fit of the appropriate MSIAH-VECM specification by calculating the ratio of the $\bar{R}^2$ and the residual variance from each estimated MSIAH-VECM to the corresponding measure for its best linear VECM counterpart. In all cases we found that the estimated MSIAH-VECM outperforms the best alternative linear VECM as this is measured by the improvement of the $\bar{R}^2$ and the reduction in the residual variance.

[17] Dacco and Satchell, (1999), Neely and Sarno (2002), Rapach and Wohar (2006) provide evidence which are in line with our arguments, although they also show that the forecasting performance of the Markov switching models do not producr good forecasts.

of interest rate differential were statistically significant which implies that, in Regime 1, these monetary fundamentals contributed most to the adjustment in restoring any deviations from the long-run equilibrium. This is consistent with the fact that during the interwar period the estimated transition probability is near or equal to unity. In addition it is clear that the exchange rate arrangements were not stable during that period. This evidence is also consistent with the adoption of fixed exchange rates under the Bretton Woods system. In Regime 2 the estimated coefficients of the exchange rate and relative money were statistically significant and thus the adjustment of equilibrium was achieved through the changes in these two variables. Again, the estimated probability of being in Regime 2 was near or equal to unity for the period during the recent period of flexible exchange rates.

For the case of Belgium we estimated an MSIAH(2)–VECM(2) model. The analysis of the estimated error correction terms shows that for the case of Regime 1 the only statistically significant coefficient is that of the exchange rate. Therefore, we argue that it is the exchange rate that adjusts to any deviations from the long-run equilibrium. This is consistent with the flexible exchange rate system that was in force up to the beginning of the interwar period. Certainly in this case we also observe a large number of switches in transition probabilities given the adoption of several alternative exchange rate arrangements. In Regime 2 we observe that the error correction coefficients of the nominal exchange rate, the relative money and interest rate differential were statistically significant and therefore all three variables adjust to bring the system to its long-run equilibrium. The probability of being in regime 1 during the post-Bretton Woods floating rate period is close to unity for the post-1979 period which also marks the establishment of the European Monetary System. When we consider the fixed exchange rate period 1944-1979 the probability of being in Regime 2 is near or equal to unity.

For the case of Canada the estimation of the switching regime model we found that the relative money and the real interest rate differential adjust to restore any deviations from the long-run equilibrium since their error correction terms were statistically significant during Regime 1. This is consistent with the gold standard period up to 1914 and the interwar period of the gold standard exchange regime from 1926 to 1933. This also holds for part of the Bretton Woods period since the early 1960s. Thus, the derived transition probabilities of being in Regime 1 are almost always equal to unity for these periods. For Regime 2 the error correction terms of exchange rate and real GDP were statistically significant which implies that they both contributed to the adjustment to the long-run equilibrium after any departure from it. This result is consistent with the fact that during the period 1920-1926 as well as during the 1950s Canada adopted a flexible exchange rate regime. Canadian dollar is also freely float currency since 1973. The estimated transition probabilities are almost always equal to unity in the case of Regime 2 for the corresponding period.

In Denmark the error correction coefficient of the exchange rate and the interest rate differential were the only statistically significant coefficients during Regime 1 and therefore these variables adjusted to restore any deviations from the long-run equilibrium. This finding is consistent with the exchange rate arrangements adopted in Denmark during the interwar period. The probability of being in Regime 1 was almost always unity during the period of flexible exchange rates. In contrast during Regime 2 the real output differential and the interest rate differential had statistically significant error correction terms and therefore these variables were the ones that provided the adjustment. This result is again consistent with the transition probabilities since it is consistent with the adoption of a fixed exchange and a target zone exchange rate regime. Overall we also noted that there were frequent changes in the regimes.

In Finland we also identified two regimes. In the case of Regime 2 the error correction coefficient of the exchange rate was statistically significant and therefore it was the variable that adjusted to restore deviations from the long-run equilibrium. This is consistent with the transition probabilities for Finland since the probability of being in Regime 2 was almost always unity during the early years of the sample when Finland's currency was under a flexible exchange rate system. In contrast, during Regime 1 it was the monetary fundamentals that adjusted since the error correction coefficient of the relative money supply and of the real output differential were statistically significant. This is consistent with the fact that Regime 1 coincided with the adoption of a fixed exchange rate regime under the Bretton Woods agreement by the Finish monetary authorities.

For France we found that in Regime 1 the error correction coefficients were statistically significant for the exchange rate and the monetary fundamentals and therefore both the exchange rate and the monetary fundamentals adjusted to deviations from long-run equilibrium. This is consistent with the gold standard prevailed until 1914 and the subsequent adoption of flexible exchange rate from 1919 to 1926 and the return to the gold standard at the end of 1926 until the end of WWII. It is also consistent with the Bretton Woods fixed exchange rate system. We also noted that during this period the exchange rate exhibited some volatility which ceased to exist after the adoption of a fixed exchange rate system. In Regime 2 the adjustment to long-run equilibrium came from the exchange rate and the relative real output differential which had the statistically significant error correction term. This is consistent with the adoption of the post-Bretton Woods flexible exchange rate regime since 1979.

For Italy the estimation of the two regimes led to the conclusion that during Regime 1 the adjustment to the long-run equilibrium came only from the monetary fundamentals. This is what one would expect during fixed exchange rates periods and is related to the estimated transition probabilities. The derived transition probabilities show that being in Regime 1 was near or equal to unity during the period of fixed exchanges rates adopted at Bretton Woods. Similar evidence is drawn for the period up until the interwar period and once again we observe a large number of regime switching. For Regime 2 it was the exchange rate and the interest rate differential that adjusted to any departures from the long-run equilibrium, which was consistent with the post-Bretton Woods floating exchange rates. This was confirmed by the transition probabilities since it is shown that during the recent flexible exchange regime these were almost always unity.

For the Netherlands we also found that during Regime 1 the exchange rate along with the real output and interest rate differential adjusted to any departures from the long-run equilibrium. The transition probabilities reveal that the probability of the Netherlands being in Regime 1 was near or equal to unity for the interwar period and the post-Bretton Woods flexible exchange rate period. As expected, the transition probabilities for the interwar period again exhibit a large number of switches. In Regime 2 we found that only the monetary fundamental variables adjust to restore deviations from long-run equilibrium in this regime. This is consistent with the estimated transition probabilities which are near or equal to unity within this regime.

For Norway we found that for Regime 1 both the relative money supplies and the real output differential adjusted to the long-run equilibrium. It is shown from the estimated transition probabilities that the probability of Norway being in the high volatility (flexible exchange rate regime) was near or equal to unity until up to the interwar period and during the Bretton Woods exchange rates system. During these periods the adjustment came from the monetary fundamentals. In Regime 2, the error correction coefficients of the exchange rate, the real output and the interest rate differential were statistically significant. Therefore, these variables adjusted to restore deviations from the long-run equilibrium under the flexible exchange rate regime that was implemented since 1973 and in addition for a short period of flexible exchange rates in the 1920s. The corresponding transition probabilities show that the probability of being in Regime 2 was near or equal to unity.

For Portugal the estimation of the two regimes led to the conclusion that during Regime 1 the adjustment to the long-run equilibrium came only from the relative money supplies and the interest rate differential. Therefore, these variables adjusted to restore deviations from the long-run equilibrium during the gold standard period, the interwar period and the fixed exchange rates system prevailed form 1944-1973. The derived transition probabilities show that being in Regime 1 was near or equal to unity during these periods of international monetary arrangements. For the period up until the interwar period and once again we observe a large number of regime switching. For Regime 2 it was the exchange rate and the interest rate differential that adjusted to any departures from the long-run equilibrium, which was consistent with the post-Bretton Woods floating exchange rates. This was confirmed by the transition probabilities since it is shown that during the recent flexible exchange regime these were almost always unity.

For the case of Spain we estimated an MSIAH(2)–VECM(3) model. The analysis of the estimated error correction terms shows that for the case of Regime 1 the only statistically significant coefficient is that of the exchange rate. Therefore, we argue that it is the exchange rate that adjusts to any deviations from the long-run equilibrium. This is consistent with the flexible exchange rate system that was in force up to the beginning of the interwar period. Certainly in this case we also observe a large number of switches in transition probabilities given the adoption of several alternative exchange rate arrangements. In Regime 2 we observe that the error correction coefficients of the nominal exchange rate, the relative money and interest rate differential were statistically significant and therefore all three variables adjust to bring the system to its long-run equilibrium. The probability of being in regime 1 during the post-Bretton Woods floating rate period is close to unity for the post-1979 period which also marks the establishment of the European Monetary System. When we consider the fixed exchange rate period 1944-1979 the probability of being in Regime 2 is near or equal to unity.

For the case of Sweden the estimation of the switching regime model we found that the relative money and the interest rate differential adjust to restore any deviations from the long-run equilibrium since their error correction terms were statistically significant during Regime 1. This is consistent with the gold standard period up to 1914 and the interwar period of the gold standard exchange regime from 1926 to 1933. This also holds for the Bretton Woods period since the early 1960s. Thus, the derived transition probabilities of being in Regime 1 are almost always equal to unity for these periods. For Regime 2 the error correction terms of exchange rate and real GDP were statistically significant which implies that they both contributed to the adjustment to the long-run equilibrium after any departure from it. The estimated transition probabilities are almost always equal to unity in the case of Regime 2 for the corresponding period.

For Switzerland in Regime 1 the error correction coefficients of relative money, real output differential and interest rate differential were statistically significant and therefore these were the variables that adjusted most to any deviations from the long-run equilibrium. This is clearly indicated that the transition probabilities in Regime 1 during the period up until the interwar period and during the Bretton Woods period (both fixed exchange rate periods) were almost always unity. In the case of Regime 2, the coefficients of the exchange rate as well of the real output and interest rate differential were statistically significant. This result is consistent with the current flexible exchange rates period and it is observed in the transition probabilities since they take a value of unity during the period of flexible exchange rates between during the specific period.

Finally, in the case of the United Kingdom the estimation of the two regimes led to the conclusion that during Regime 1 the adjustment to the long-run equilibrium came only from the relative money supplies and the interest rate differential. Therefore, these variables adjusted to restore deviations from the long-run equilibrium during the gold standard period, the interwar period and the fixed exchange rates system prevailed form 1944-1973. The derived transition probabilities show that being in Regime 1 was near or equal to unity during these periods of international monetary arrangements. For the period up until the interwar period and once again we observe a large number of regime switching. For Regime 2 it was the exchange rate and the interest rate differential that adjusted to any departures from the long-run equilibrium, which was consistent with the post-Bretton Woods floating exchange rates. This was confirmed by the transition probabilities since it is shown that during the recent flexible exchange regime these were almost always unity.

In order to assess the regime qualification performance of the chosen Markov-switching models, we calculated the Regime Classification Statistic (RCM) developed by Ang and Bekaert (2002a,b). This measure is based on the fact that the ex-post (smoothed) probabilities p_t are close either to one or zero and therefore a good regime-switching model should classify regimes sharply. The RCM for a model with two regimes may be calculated as follows:

$$RCM(2) = 400 * \frac{1}{T}\sum_{i=1}^{T} p_t(1-p_t) \qquad (21)$$

where T is the sample size, p_1 and $(1-p_1)$ is the smoothed probability of being in regime $j = 1,2$ at time t, and RCM takes values between zero and one hundred. In general the lower the value of RCM the better the performance of the model is. The ideal model will have an RCM with a value close to zero. Weak regime inference implies that regime-switching models cannot distinguish among regimes based on the behaviour of data and this may be due to misspecification. A model which poorly distinguishes between regimes will have an RCM with a value close to 100. Table 6 reports the calculated RCM statistic for the full sample which is close to zero for all countries, implying a very satisfactory regime classification.

6. Summary and Concluding Remarks

In this paper we provide an extended review of the monetary model of exchange rate over the last forty years. Furthermore, we provide further study of the monetary model under regime switching. We analyzed the case of the link between exchange rates and monetary fundamentals for fourteen industrialized countries. We considered the presence of nonlinearities in the relationship between the nominal bilateral exchange rate and macroeconomic fundamentals and we estimated the appropriate Markov-switching vector error correction model. We used annual data spanning from the late nineteenth century or early twentieth century to the late twentieth century for the bilateral nominal exchange rates against the US dollar and the respective macroeconomic fundamentals.

Furthermore, from a methodological point of view it was important to examine the adjustment mechanisms to the long-run equilibrium. Since the period under examination covers a number of alternative international monetary arrangements, such the gold standard, the Bretton Woods period and the current flexible exchange rate regime, our analysis focused on demonstrating whether the exchange rate or the macroeconomic variables were the main vehicle in achieving their target. Therefore, it was important to reveal whether the adjustment back to equilibrium took place primarily through the nominal exchange rate during periods of floating exchange rates and through monetary fundamentals during periods in which some variant of a fixed exchange rate system was in force. To this end the monetary model provided the appropriate framework to study the behaviour of exchange rate movements in periods of transition.

There are several important findings that stem from the present analysis. First for each bilateral exchange rate and the respective macroeconomic variables we were able to capture nonlinearities with the estimation of the appropriate Markov switching regime model with two regimes. The fitted model was quite general since it allowed for regime shifts in the intercept and the complete set of parameters, as well as the variance-covariance matrix. In addition, for all cases the null hypothesis of linearity was rejected when tested against the alternative of a MS-VECM specification. Second, our analysis has clearly shown that during the period when some variant of fixed exchange rates was adopted in each country, the monetary fundamentals adjust to restore deviations from the long-run equilibrium. In contrast during periods with less restricted exchange rate regimes, it was the exchange rate that adjusted to restore any disequilibrium. Finally, the application of the Regime Classification Measure developed by Ang and Bekaert (2002a, b) showed that our estimated Markov-switching models distinguished very well between the two regimes.

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Table 1: Unit root and stationarity tests

Variable	ADF t-tests t_μ	t_τ	DF-GLS$_u$ t_μ	t_τ	MZ_a^{GLS}	MZ_t^{GLS}	KPSS-tests η_μ	η_τ
Australia								
e	-0.471	-1.408	-1.301	-2.033	-5.671	-2.288*	1.921*	0.262*
$m-m^*$	-0.507	-1.885	-1.623	-2.083	1.299	2.001	2.033*	0.366*
$y-y^*$	-1.103	-1.277	-1.812	-1.129	-3.213	-2.061*	1.904*	0.152*
$\iota-\iota^*$	-0.905	-1.566	-2.651	-1.356	-4.889	-2.886*	0.987*	0.205*
Belgium								
e	1.032	-1.075	0.792	-0.454	-7.566	1.067	1.546*	0.513*
$m-m^*$	0.056	-1.596	0.367	-1.103	1.098	-3.233	1.335*	0.292*
$y-y^*$	-0.312	-1.723	-0.385	-1.044	-9.051*	-1.045	2.155*	0.441*
$\iota-\iota^*$	-0.225	-2.191	-2.891*	-3.011	-6.667	-1.889	1.998*	0.393*
Canada								
e	-0.151	-1.603	1.956	-1.233	-8.002	-1.776	0.887*	0.311*
$m-m^*$	1.133	-2.276	0.445	0.088	-5.787	1.093	0.509*	0.218*
$y-y^*$	-0.335	-3.998*	-0.487	-0.177	-5.576	0.098	0.443*	0.191*
$\iota-\iota^*$	-1.156	-0.098	-1.766	1.089	-9.206*	0.205	1.019*	0.166*
Denmark								
e	-0.201	-1.558	2.023	-1.245	2.556.	-0.991	0.355*	0.120
$m-m^*$	-1.225	-4.228*	0.306	0.091	-9.776*	0.097	0.322	0.303*
$y-y^*$	-0.335	-3.998*	-0.398	-3.445*.	-2.226	-0.998	0.591*	0.202*
$\iota-\iota^*$	-1.155	-2.001	-0.721	1.331	-4.121	-1.445	0.519*	0.291*
Finland								
e	-0.151	-1.603	1.956	-1.233	-8.002	-1.776	0.887*	0.311*
$m-m^*$	1.133	-2.276	0.445	0.088	-5.787	1.093	0.509*	0.218*
$y-y^*$	-0.335	-3.998*	-0.487	-0.177	-5.576	0.098	0.443*	0.191*
$\iota-\iota^*$	-1.156	-0.098	-1.766	1.089	-9.206*	0.205	1.019*	0.166*
France								
e	-0.344	-1.901	-1.001	-1.036	-2.361	-1.445	0.503*	0.256*
$m-m^*$	1.209	-1.302	-1.122	-1.265	-3.678	-1.887	0.398*	0.181
$y-y^*$	-0.147	-1.335	-0.609	-0.305	-9.305*	-1.023	0.201	0.301*
$\iota-\iota^*$	-1.304	-0.101	-2.608	-2.645	2.001	1.103	0.491*	0.105

Italy								
e	-1.902	-1.701	-2.132	-2.098*	-6.222	-1.332	0.698*	0.115
$m-m^*$	-1.335	-2.444	-1.228	0.088	1.116	1.220	0.672*	0.318*
$y-y^*$	-0.701	-1.299	-1.187	-0.662	1.105	0.122	0.433	0.202*
$\iota-\iota^*$	-1.338	-1.409	-1.609	-1.116	-6.225	-2.445*	0.672*	0.208*
Netherlands								
e	-0.300	-1.558	-1.336	-0.609	-9.044*	-1.446	0.498*	0.177*
$m-m^*$	-0.892	-3.405*	-1.288	-0.293	1.667	-1.111	0.508*	0.112
$y-y^*$	-1.208	-2.988*	-0.897	-0.209	-7.889	-1.209	0.307	0.299*
$\iota-\iota^*$	-1.156	-0.098	-1.333	-1.307	-5.332	1.885	0.625*	0.191*
Norway								
e	-1.988	-1.206	-0.445	-2.333	2.063	0.898	0.661*	0.133
$m-m^*$	-0.998	-1.665	-2.002	-1.889	1.099	1.693	0.701*	0.155*
$y-y^*$	-1.556	-2.003	-2.113	-1.233	-4.433	0.909	0.551*	0.177*
$\iota-\iota^*$	-1.200	-0.552	-1.103	-2.092	-7.229*	0.805	0.908*	0.206*
Portugal								
e	-0.909	-3.243*	-1.990	-0.998	-6.443	-1.223	0.323*	0.221*
$m-m^*$	-1.561	-1.089	-1.065	0.902	-5.609	-2.001	1.202*	0.167*
$y-y^*$	-1.612	-1.345	-1.233	-1.361	0.999	-1.022	0.399	0.222*
$\iota-\iota^*$	-0.901	-1.223	-1.612	0.803	-11.225*	0.667	0.554*	0.111
Spain								
e	-0.113	-1.655	-1.361	-1.208	-3.223	-2.077*	0.498*	0.221*
$m-m^*$	-1.215	-1.099	-1.108	-1.100	-3.998	-3.889*	0.737*	0.299*
$y-y^*$	-0.909	-1.244	-0.665	-0.882	-10.28*	-1.756	0.692*	0.401*
$\iota-\iota^*$	-2.021	-1.344	-1.335	-0.998	-7.355	-1.566	1.203*	0.303*
Sweden								
e	-1.209	-1.445	-1.225	-0.882	-4.772	0.995	0.499*	0.166*
$m-m^*$	-1.335	-1.003	0.668	-0.904	-2.668	1.701	0.509*	0.218*
$y-y^*$	-0.335	-3.998*	-0.487	-0.177	-1.989	-2.001	0.771*	0.227*
$\iota-\iota^*$	-1.334	-1.345	-3.781*	-1.267*	-6.332	0.309	0.899*	0.321*
Switzerland								
e	-1.309	-1.771	-1.442	-1.609	-3.225	0.881	0.883*	0.224*
$m-m^*$	1.133	-2.276	0.901	0.771	-5.628	-1.223	0.495*	0.333*
$y-y^*$	-1.613	-2.991*	-2.551	-2.991	-3.229	2.223	0.513*	0.205*
$\iota-\iota^*$	-0.992	-0.598	-1.334	1.066	-5.228	-2.281*	0.819*	0.332*

United Kingdom

e	-0.891	-2.881*	1.001	-1.281	-3.818	-1.229	0.332*	0.161*
$m-m^*$	-0.391	-1.981	0.901	0.209	-3.776	1.093	0.509*	0.218*
$y-y^*$	-0.335	-3.998*	-0.487	-0.177	-5.576	0.098	0.443*	0.191*
$\iota-\iota^*$	-1.156	-0.098	-1.766	1.089	-9.206*	0.205	1.019*	0.166*

Notes: e, $m-m^*$, $y-y^*$, $\iota-\iota^*$ are the nominal exchange rate, relative money supply and relative real output respectively nominal interest rate differential.

- t_μ and t_τ are the standard augmented Dickey-Fuller test statistics when the relevant auxiliary regression contains a constant and a constant and a trend respectively. The number of lagged differenced terms required for serial correlation correction in the ADF auxiliary regressions is selected on the basis of a general to specific testing strategy which is terminated when a sequence of t-ratio elimination tests on the lagged differenced terms leads to a rejection at the 10% significance level and the residuals of the resultant specification satisfy standard misspecification testing (Perron and Ng, 1996). The response surface regressions of MacKinnon (1991, 1994) are used for determining the significance of the ADF test statistics.
- The DF-GLS_u by Elliott (1999) is a test with an unconditional alternative hypothesis. The critical values for the DF-GLS_u test at the 1%,5% and 10% significance level are: -3.28, -2.73, -2.46 (with constant) and –3.71,-3.17, -2.91 (with constant and trend), respectively (Elliott,1999).
- MZ_α and MZ_t are the Ng and Perron (2001) GLS versions of the Phillips-Perron tests. The critical values at 5% significance level are: -8.10 and -1.98 (with constant and with constant and trend), respectively (Ng and Perron, 2001, Table 1).
- η_μ and η_τ are the KPSS test statistics for level and trend stationarity respectively (Kwiatkowski *et al.,* 1992). For the computation of theses statistics a Newey and West (1994) robust kernel estimate of the "long-run" variance is used. The kernel estimator is constructed using a quadratic spectral kernel with VAR(l) pre-whitenning and automatic data-dependent bandwidth selection [see, Newey and West, 1994 for details]. The 5% critical values for level and trend stationarity are 0.461 and 0.148 respectiveley, and they are taken from Sephton (1995, Table 2).

* and ** indicate significance at the 95% and 99% confidence level respectively.

Table 2: Structural breaks and nonlinear unit root tests

Variable	Zivot-Andrews		MZ_a^{GLS}	MZ_t^{GLS}	t_{NL}^c	t_{NL}^t
	t_μ	t_τ				
<u>Australia</u>						
e	-3.938	-3.639	-28.763	-2.831	-2.431	-2.454
$m-m^*$	-3.405	-2.223	-19.012	-2.651	-3.001	-3.445
$y-y^*$	-4.114	-4.278	-28.315	-2.562	-2.053	-3.001
$\iota-\iota^*$	-1.109	-1.925	-27.251	-1.983	-2.981	-3.521
<u>Belgium</u>						
e	-3.994	-3.665	-26.332	-2.032	-2.852	-3.212
$m-m^*$	-2.226	-1.453	-22.331	-3.051	-3.032	-3.578
$y-y^*$	-2.998	-2.667	-28.219	-2.668	-1.988	-2.013
$\iota-\iota^*$	-1.688	-3.911	-27.665	-2.883	-2.033	-1.987
<u>Canada</u>						
e	-3.669	-2.998	-29.336	-3.122	-1.677	-2.988
$m-m^*$	-4.002	-4.233	-26.665	-2.881	-1.979	-2.556
$y-y^*$	-2.338	-4.227	-28.334	-3.229	-1.455	-3.292
$\iota-\iota^*$	-3.771	-2.135	-27.887	-3.166	-2.199	-2.989
<u>Denmark</u>						
e	-2.013	-2.277	-25.344	-2.155	-2.332	-2.575
$m-m^*$	-3.504	-3.609	-23.434	-2.681	-1.808	-2.001
$y-y^*$	-2.592	-2.698	-27.557	-3.105	-1.809	-3.168
$\iota-\iota^*$	-3.202	-2.168	-30.252	-3.200	-2.509	-1.628
<u>Finland</u>						
e	-2.108	-2.206	-26.107	-2.771	-1.803	-2.441
$m-m^*$	-4.112	-4.233	-28.105	-4.065*	-2.661	-3.701*
$y-y^*$	-4.651*	-3.105	-29.102	-3.606	-1.901	-4.288*
$\iota-\iota^*$	-3.223	-2.991	-28.993	-3.221	-2.709	-2.665
<u>France</u>						
e	-3.005	-2.103	-24.391	-2.189	-2.558	-2.228
$m-m^*$	-3.661	-4.557	-29.883	-3.477	-3.051*	-2.705
$y-y^*$	-5.031*	-4.512	-24.205	-3.688	-2.099	-3.105
$\iota-\iota^*$	-4.221	-2.908	-23.689	-2.704	-2.101	-2.609

Italy						
e	-3.208	-2.243	-28.344	-3.208	-2.402	-2.669
$m-m^*$	-3.789	-4.198	-29.228	-2.667	-1.109	-2.109
$y-y^*$	-2.501	-4.099	-30.500	-2.990	-2.661	-3.701*
$\iota-\iota^*$	-2.990	-2.301	-29.809	-3.202	-2.509	-3.001
Netherlands						
e	-3.558	-2.701	-24.220	-3.101	-2.755	-3.209
$m-m^*$	-4.002	-4.233	-26.665	-2.881	-1.979	-2.556
$y-y^*$	-2.702	-2.621	-21.488	-2.991	-1.728	-3.552
$\iota-\iota^*$	-3.609	-4.235	-30.702	-3.555	-2.402	-2.559
Norway						
e	-3.911	-2.776	-27.209	-3.403	-1.902	-2.333
$m-m^*$	-3.766	-2.244	-29.883	-2.454	-2.023	-3.005
$y-y^*$	-2.523	-1.992	-29.300	-3.709	-1.989	-3.902
$\iota-\iota^*$	-3.200	-2.900	-30.505	-3.208	-2.098	-2.009
Portugal						
e	-3.928	-3.111	-28.554	-3.333	-2.012	-3.566
$m-m^*$	-4.013	-4.544	-27.883	-3.026	-2.171	-2.388
$y-y^*$	-2.905	-4.305	-30.922	-3.441	-1.988	-3.609
$\iota-\iota^*$	-4.238	-4.501	-28.330	-2.202	-2.409	-2.108
Spain						
e	-3.222	-2.446	-29.708	-3.601	-1.901	-2.101
$m-m^*$	-4.109	-4.207	-29.323	-2.701	-2.334	-2.198
$y-y^*$	-1.997	-1.825	-21.201	-3.662	-2.233	-2.779
$\iota-\iota^*$	-3.106	-2.442	-27.228	-3.487	-2.601	-2.101
Sweden						
e	-3.441	-2.335	-30.202	-3.559	-1.773	-2.619
$m-m^*$	-3.334	-4.018	-29.305	-2.440	-2.778	-3.021
$y-y^*$	-2.779	-4.232	-29.445	-3.338	-2.988	-3.115
$\iota-\iota^*$	-3.556	-2.991	-30.669	-3.203	-2.443	-2.601
Switzerland						
e	-3.258	-2.781	-26.446	-3.337	-2.708	-3.202
$m-m^*$	-3.801	-4.509	-23.778	-3.232	-2.889	-2.309
$y-y^*$	-4.555	-4.908	-29.406	-3.454	-1.991	-2.404
$\iota-\iota^*$	-2.881	-2.709	-29.171	-3.209	-2.206	-2.508

United Kingdom

e	-3.202	-2.504	-29.552	-3.773	-1.388	-2.052
$m-m^*$	-3.776	-4.198	-27.556	-2.768	-1.882	-2.602
$y-y^*$	-3.199	-2.661	-27.558	-3.102	-1.669	-2.882
$\iota-\iota^*$	-3.192	-2.347	-30.199	-3.637	-2.500	-3.400

Notes: e, $m-m^*$, $y-y^*$, $\iota-\iota^*$ are the nominal exchange rate, relative money supply and relative real output respectively nominal interest rate differential.

- t_{μ} and t_{τ} are Zivot and Andrews test statistics for the null hypothesis of a unit root against the alternative of stationarity with a structural break. The critical values at 5% significance level are: -4.80 and -5.08 (with constant and with constant and trend), respectively (Zivot and Andrews, 1992, Table 1).
- MZ_{a}^{GLS} and MZ_{t}^{GLS} are the Perron and Rodriguez (2003) GLS versions to the case where a change in the trend function is allowed to occur . The critical values at 5% significance level are: -31.04 and -3.91 (with constant and with constant and trend), respectively (Perron and Rodriguez, 2003, Table 1).
- t_{NL}^{c} and t_{NL}^{t} are the nonlinear unit root tests with the null of nonstationarity against the alternative of a nonlinear exponential smooth transition autoregressive (ESTAR) process, with a constant and a constant and a linear trend, respectively. The asymptotic critical values at the 5% (1%) critical values for an equation with a constant and a constant and a linear trend are -2.93 (-3.48) and -3.40 (-3.93) respectively, (Kapetanios *et al*., 2003, Table 1).

  * and ** indicate significance at the 95% and 99% confidence level respectively.

Table 3: Johansen-Juselius cointegration trace tests results

$$e_t = \beta_0 + \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t - i_t^*)$$

Country	**(n-r)**	**r**
Australia	3	58.93*
	2	18.88
	1	7.42
Belgium	3	67.33*
	2	17.34
	1	8.81
Canada	3	61.12*
	2	16.23
	1	6.33
Denmark	3	45.21*
	2	22.44
	1	7.86
Finland	3	46.38*
	2	19.02
	1	7.13
France	3	39.12*
	2	17.90
	1	5.18
Italy	3	40.02*
	2	17.22
	1	4.14
Netherlands	3	41.13*
	2	18.02
	1	7.63
Norway	3	51.15*
	2	17.96
	1	7.15
Portugal	3	55.16*
	2	18.35
	1	6.98

Spain	3	44.12*
	2	17.65
	1	7.23
Sweden	3	42.16*
	2	18.33
	1	8.19
Switzerland	3	53.16*
	2	18.32
	1	8.35
United Kingdom	3	50.12*
	2	18.86
	1	7.87

Notes: r denotes the number of eigenvectors and (n-r) ia the number of common trends. Trace is the Johansen Trace likelihood ratio statistic. A structure of four lags was chosen according to a likelihood ratio test, corrected for the degrees of freedom (Sims, 1980) and the Ljung-Box Q statistic for detecting serial correlation in the residuals of the equations of the VAR. A model with a constant restricted in the cointegrating vector is chosen according the Johansen (1992) testing strategy. (*) denotes statistical significance at the five percent critical level. The 5% critical values are 34.91, 19.96 and 9.24 respectively and they are taken from MacKinnon *et al*. (1999, Table IV).

Table 4: "bottom-up" procedure for model specification

Model	LR1 – *p-value*[*]	LR2 - $x^2_{(g)}$	LR3 - $x^2_{(g)}$	AIC Ratio	HQ Ratio	SIC Ratio
Australia	441.56 (0.0000)	168.22 ($x^2_{(10)} = 18.3$)	891.23 ($x^2_{(196)} = 229.6$)	1.1126	1.0459	0.9286
Belgium	333.220 (0.0000)	155.12 ($x^2_{(10)} = 18.3$)	334.22 ($x^2_{(36)} = 50.99$)	1.0531	1.0390	1.0174
Canada	541.776 (0.0000)	180.13 ($x^2_{(10)} = 18.3$)	609.22 ($x^2_{(126)} = 153.1$)	1.1314	1.0961	1.0390
Denmark	357.111 (0.0000)	201.56 ($x^2_{(10)} = 18.3$)	443.28 ($x^2_{(84)} = 106.3$)	1.0626	1.0210	0.9530
Finland	507.566 (0.0000)	887.34 ($x^2_{(10)} = 18.3$)	77.89 ($x^2_{(36)} = 50.99$)	1.0980	1.0489	0.9667
France	246.212 (0.0000)	367.28 ($x^2_{(10)} = 18.3$)	272.06 ($x^2_{(84)} = 106.3$)	1.0088	0.9797	0.9337
Italy	271.158 (0.0000)	303.66 ($x^2_{(10)} = 18.3$)	256.28 ($x^2_{(84)} = 106.3$)	1.0002	0.9708	0.9243
Netherlands	528.022 (0.0000)	121.90 ($x^2_{(10)} = 18.3$)	298.23 ($x^2_{(84)} = 106.3$)	1.1858	1.1382	1.0580
Norway	293.793 (0.0000)	180.67 ($x^2_{(10)} = 18.3$)	129.65 ($x^2_{(52)} = 69.83$)	1.0453	1.0208	0.9826
Portugal	317.332 (0.0000)	233.55 ($x^2_{(10)} = 18.3$)	168.33 ($x^2_{(68)} = 88.25$)	1.0463	1.0154	0.9664
Spain	521.335 (0.0000)	356.13 ($x^2_{(10)} = 18.3$)	157.12 ($x^2_{(52)} = 69.83$)	1.0566	1.0233	0.9732
Sweden	489.221 (0.0000)	602.51 ($x^2_{(10)} = 18.3$)	150.13 ($x^2_{(68)} = 88.25$)	1.0998	1.0245	0.9891
Switzerland	502.113 (0.0000)	589.23 ($x^2_{(10)} = 18.3$)	373.28 ($x^2_{(84)} = 106.3$)	1.1467	1.0325	0.9624
United Kingdom	494.223 (0.0000)	433.28 ($x^2_{(10)} = 18.3$)	201.17 ($x^2_{(68)} = 88.25$)	1.0858	1.1258	0.9765

Note: The p-values are given based on the likelihood ratio test (LR) for the null of a linear VECM. The value in squared brackets next to LR is the marginal significance level of this test, based on Davies (1987). LR1 tests the null hypothesis that there is no regime switching. LR2 tests the null hypothesis that there is no regime switching in the autoregressive parameters and in the variance-covariance matrix [i.e. MSI (2)-VECM(p) against MSIH(2)-VECM(p)]. LR3 tests the null hypothesis that there is no regime switching in the autoregressive parameters [i.e. MSIH(2)-VECM(p) against MSIAH(2)-VECM(p)]. The statistical criteria LR2 and LR3 are distributed as a x^2 with g degrees of freedom, where g is the number of restrictions. AIC, HQ and SIC denote the Akaike information criterion, Hannan Quinn and Schwartz information criterion respectively, and they provide the ratios between the chosen MS-VECM model and the respective linear VECM model.

Table 5. Regime dependent error correction terms

$\alpha_{\Delta s_t}(z=1)$	$\alpha_{\Delta s_t}(z=2)$	$\alpha_{\Delta m_t}(z=1)$	$\alpha_{\Delta m_t}(z=2)$	$\alpha_{\Delta y_t}(z=1)$	$\alpha_{\Delta y_t}(z=2)$	$\alpha_{\Delta i_t}(z=1)$	$\alpha_{\Delta i_t}(z=2)$
Australia– MSIAH(2)–VECM(5)							
0.0010 (0.007)	0.221* (0.058)	-0.535* (0.265)	-0.106* (0.051)	0.441 (0.597)	-0.033 (0.076)	0.155* (0.046)	0.112 (0.215)
Belgium – MSIAH(2)–VECM(2)							
-0.063* (0.004)	-0.071* (0.005)	-0.004 (0.002)	-0.056* (0.002)	0.007 (0.008)	0.000 (0.025)	0.010 (0.006)	-0.022* (0.001)
Canada – MSIAH(2)–VECM(7)							
-0.002 (0.001)	-0.017* (0.001)	-0.104* (0.011)	-0.006 (0.012)	-0.010 (0.024)	0.324* (0.058)	-0.021* (0.004)	0.003 (0.007)
Denmark – MSIAH(2)–VECM(5)							
-0.054* (0.012)	-0.012 (0.033)	0.007 (0.010)	0.006 (0.021)	-0.045 (0.056)	-0.071* (0.022)	-0.022* (0.004)	-0.058* (0.012)
Finland– MSIAH(2)–VECM(5)							
-0.022 (0.030)	0.091* (0.022)	-0.903* (0.191)	-0.021 (0.063)	0.298* (0.074)	0.168 (0.065)	0.031 (0.019)	0.101 (0.093)
France – MSIAH(2)–VECM(6)							
0.088* (0.025)	0.093* (0.012)	-0.189* (0.033)	-0.022 (0.081)	0.303* (0.098)	-0.405* (0.111)	0.196* (0.034)	0.008 (0.023)
Italy – MSIAH(2)–VECM(5)							
-0.022 (0.071)	-0.021 (0.044)	0.057* (0.010)	0.007 (0.021)	-0.091* (0.029)	-0.004 (0.012)	0.103* (0.018)	-0.005 (0.023)
Netherlands – MSIAH(2)–VECM(5)							
-0.098* (0.015)	-0.012 (0.025)	0.041 (0.097)	-0.202* (0.051)	-0.052* (0.023)	-0.109* (0.032)	-0.219* (0.035)	0.065* (0.019)
Norway – MSIAH(2)–VECM(3)							
-0.008 (0.021)	-0.077* (0.013)	-0.066* (0.018)	0.011 (0.023)	0.151* (0.032)	-0.037* (0.08)	0.007 (0.006)	0.053* (0.009)
Portugal – MSIH(2)–VECM(4)							
0.003 (0.012)	-0.028* (0.005)	0.067* (0.013)	0.065 (0.051)	0.014 (0.010)	0.005 (0.010)	-0.031* (0.006)	-0.105* (0.033)
Spain – MSIH(2)–VECM(3)							
0.702* (0.187)	-0.201* (0.045)	0.008 (0.013)	0.076* (0.019)	0.009 (0.021)	0.007 (0.019)	-0.002 (0.007)	-0.201* (0.065)
Sweden – MSIH(2)–VECM(4)							
0.021 (0.081)	-0.072* (0.017)	0.053* (0.009)	0.017 (0.021)	0.008 (0.016)	0.541* (0.105)	-0.443* (0.098)	-0.023 (0.054)
Switzerland – MSIH(2)–VECM(7)							
0.008 (0.027)	-0.021* (0.006)	0.171* (0.041)	0.027 (0.031)	0.202* (0.034)	0.092* (0.021)	-0.031 (0.054)	-0.281* (0.077)
United Kingdom – MSIH(2)–VECM(5)							
0.015 (0.031)	-0.206* (0.044)	0.071* (0.016)	0.059 (0.078)	0.019 (0.021)	0.011 (0.039)	-0.077* (0.013)	-0.322* (0.055)

Note: $\alpha_{\Delta s_t}$, $\alpha_{\Delta m_t}$, $\alpha_{\Delta y_t}$ and $\alpha_{\Delta i_t}$ denote the estimated coefficients of the error correction terms for the exchange rate, the money supply differential, the output differential and the interest rates differential in both regimes 1 and 2. Figures in parentheses are asymptotic standard errors and (*) denotes statistical significance at the 5% level.

Table 6. Regime classification measure

Country	RCM
Australia	3.98
Belgium	5.33
Canada	1.10
Denmark	6.38
Finland	7.01
France	13.78
Italy	3.48
Netherlands	1.15
Norway	11.28
Portugal	6.92
Spain	7.02
Sweden	0.13
Switzerland	0.28
United Kingdom	2.69

Note: The statistics for a model with two regimes are calculated as $RCM(2) = 400 * \frac{1}{T}\sum_{i=1}^{T} p_t(1-p_t)$ where T is the sample size, p_1 and $(1-p_1)$ is the smoothed probability to be in regime $j = 1,2$ at time t, and *RCM* takes values between zero and one hundred. A value close to zero implies a very good discrimination between the two regimes whereas a value close to one hundred implies a model that poorly distinguishes between the two regimes.