

How Well Do Monetary Fundamentals Forecast Exchange Rates in Developing Economies?

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Abstract: The main objective of this study is to re-investigate the exchange rates predictability puzzle using the monetary model. It is hypothesised that the performance of exchange rate predictability is better off in countries with monetary instability. We employed bootstrap technique as proposed by Kilian (1999) to alleviate statistical inference intricacies inherent in the long horizon forecasting for three different monetary models (flexible price, sticky price and relative price) for selected developing economies. The empirical result shows the superiority of sticky price model along with the evidence of exchange rate predictability for high inflation economies.

Keywords: Foreign exchange, forecasting and simulation, international finance

JEL classification: F31 and F37

1. Introduction

The aim of this paper is to investigate the exchange rate forecastability puzzle that suggests that macroeconomic fundamentals contain a negligible predictive content about the movements of nominal exchange rates. Since the seminal papers by Meese and Rogoff (1983a; 1983b), much resources have been channelled into the refinement of theoretical models and advancement of estimation techniques to explain better the puzzle. However, empirical evidence from mature economies has consistently failed to overturn this paradox. Consequently, clarifying the puzzle remains a challenging area for researchers.

In this paper, we give monetary models another chance and investigate whether using dataset from developing economies can improve their forecasting performance. Our expectation is to find significant exchange rate predictability for countries with unstable macroeconomic fundamentals (see for example McNown and Wallace 1994; Rogoff 1996; 1999a; Moosa 2000). The reason underlying this hypothesis is that countries with greater monetary instability are expected to show a stronger correlation between exchange rates and monetary fundamentals. Rogoff (1999a) argues that economically stable countries like United States, Germany and Japan generally experience very modest inflation rates. In such circumstances, it is difficult to identify the effect of monetary shocks on exchange rates. On the other hand, developing economies experience high inflation rates, trade balance deficit, budget deficit and excess money supply.¹ These relatively weak economic fundamentals, in addition to the poor management of the economy, are postulated to be crucial in predicting

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¹ Refer to Table 1 for comparison between income volatility and inflation rate between developing countries and the US. The countries of Chile, Israel and Uruguay are categorised as high inflation countries.

exchange rates under the monetary approach. Furthermore, most of the literature in the area of exchange rate predictability deals with developed and industrialised economies. Until now, not much work has been done to investigate the forecastability of exchange rates in developing economies despite their increasingly liberalised financial markets and their growing importance in the global financial system.²

This study differs from most previous studies in a number of ways. First, our sample is limited to developing countries that satisfy two important assumptions of the exchange rate determination model: relatively floating exchange rate and a considerably open economy for a long period to allow meaningful time series analysis. It does not mean that the developing countries that we chose are fully liberalised, rather that the markets are satisfactorily open with little market frictions and government intervention. The countries we considered are Chile, Uruguay, Philippines, Thailand, Israel, Morocco, South Africa and Tunisia. According to Levy-Yeyati and Sturzenegger (2003) these are countries that are adopting a relatively floating exchange rate regime and are in the process of liberalising their capital account.

Second, motivated by Chinn and Meese (1995), we calculated the deviation from monetary fundamentals that are suitable for the developing economies. In particular, we considered sticky price and relative price Balassa-Samuelson monetary models to account for developing country characteristics, as suggested by MacDonald and Ricci (2001). These models are expected to be superior to the standard flexible monetary model especially for countries which are still in the process of the liberalisation period (see Crespo-Cuaresma *et al.* 2005; Candelon *et al.* 2007).

Third, we used an error-correction framework to investigate both in-sample predictive content and out-of-sample point forecast accuracy of the fundamental-based models by employing the bootstrap technique proposed by Kilian (1999). The technique is able to account for small sample biases and size distortion that arise in the inferences procedure. Furthermore, the methodology is designed to differentiate whether forecastability power (if any) is due to the contribution of the explanatory variables or simply due to the drift term in the model.

The plan of the paper is as follows. Section 2 delves into literature reviews. In Section 3, we describe the process of constructing the fundamental variables, the dataset and the econometric procedure for testing predictability of exchange rate using the monetary models. Section 4 discusses the findings and the link between predictability and economic fundamentals of developing economies. Section 5 concludes.

2. Literature Review

The study of exchange rates predictability was pioneered by Meese and Rogoff (1983a; 1983b). Their results suggest that none of the structural exchange rate models are able to

² The bulk of related studies in the developing countries are more concerned with the subject other than forecasting exchange rate movements using a monetary model. Among other issues of interest are optimal exchange rate regime, (Hochreiter and Tavlas 2004; Alfaro 2005), exchange markets integration, (Francis *et al.* 2002; Cheung *et al.* 2006; Rogers 2007; Tai 2007), exchange markets and financial crisis, (Phengpis 2006; Kan and Andreosso-O'Callaghan 2007) and exchange rate determinants (Civcir 2004; Candelon *et al.* 2006).

forecast out-of-sample better than a naïve random walk model. Subsequently, extensive work has been carried out using various econometric techniques and different information sets to challenge the superiority of the random walk over monetary models of exchange rate determination. However, after more than two decades of efforts, none of the out-of-sample empirical work finds consistent evidence of superior forecastability of structural models compared to the random walk.

Mark (1995) has given a new hope for exchange rate predictability by exploiting the assumed long-run linkages between exchange rates and monetary fundamentals. He found significant evidence of forecastability at longer horizons (12 and 16 quarter). The same conclusion can also be found in Chinn and Meese (1995) who investigated the same issue using a larger set of explanatory variables. Chinn and Meese (1995) found that fundamental-based error-correction models outperform the random walk model for long term prediction horizons. However, both the econometric techniques and the results of Mark (1995) and Chinn and Meese (1995) have not been free from criticism. Kilian (1999) found that Mark's results suffer from inconsistencies in the testing procedure and small-sample bias. Correcting for these drawbacks, Kilian (1999) found no support for long run predictability of exchange rate. Later, Berkowitz and Giorgianni (2001) argued that the results of Mark (1995) were not robust and depend heavily on the assumption of cointegration in the long run series. Berkowitz and Giorgianni (2001) showed that using the same dataset as Mark (1995) but under the unrestricted VAR model produces very little evidence of predictability. Therefore, unpredictability of exchange rates remains if no prior assumption is imposed.³

Recent studies that use different information set and econometrics approaches (mostly depart from the traditional linear time series) to analyse the association of exchange rates and economic fundamentals do find encouraging support. For example, Kilian and Taylor (2003) used an Exponential Smoothing Threshold Autoregressive (ESTAR) model for seven OECD countries. They showed the (in-sample) relevance of non linearities in exchange rate dynamics at the one- and two-year horizons. However, they still could not find support for out-of-sample predictability. Manzan and Westerhoff (2007) proposed a chartist-fundamentalist model which allows for non linear time variation in chartists' extrapolation rate that provides support for the long-term predictability for five major currencies (German mark, Japanese yen, British pound, French franc and Canadian dollar) against the US dollar. Their study shows that the fundamentalist, together with the chartist, are correcting the deviation of exchange rate from its long run equilibrium path.

Faust *et al.* (2003) criticised the use of revised data for the fundamental variables and proposed the use of real-time (unrevised) data. They argue that revised data can be used only if economic agents have the ability to predict future data (including the revision) correctly. However, this is not the case as Faust *et al.* (2005), among others, have shown that revisions to preliminary fundamental values are large and are unpredictable for some countries. Faust *et al.* (2003) empirically showed that the exchange rate determination models that use real-time data are capable of explaining about 75% of the monthly directional changes of the US dollar-Euro exchange rate.

Comprehensive debate on the reliability of long-term exchange rate forecast can be found in Berben and van Dijk (1998), Groen (1999) and Rossi (2005), among others.

A comprehensive review of the empirical literature on the exchange rate unpredictability for industrialised nations over the last few decades can be found in Neely and Sarno (2002) and Cheung *et al.* (2005). The plausible explanations for the empirical failure of the exchange rate determination models include the instability of the parameters over the period, simultaneity problems, improper modelling of expectations formation and the failure of law of one price, among others. Following these dismal findings, exchange rate economists have drawn the conclusion that exchange rate movements cannot possibly be attributed to macroeconomic fundamentals, at least in the short term. However, they have a firm belief that exchange rates cannot move independently from macroeconomic fundamentals over long horizons.

3. Exchange Rate Predictability and Developing Countries

3.1 Evidence

Investigating the predictability of exchange rate movements using a exchange rate determination model in developing economies has not been an easy task. Empirical attempts are hampered by the difficulty of finding an appropriate market that satisfies the assumption of a free floating regime, free capital mobility and a stable monetary regime.⁴ Consequently, there is only relatively little empirical evidence of exchange rate forecastability in developing countries during post-liberalisation eras. These handful of empirical studies also produced inconsistent results and therefore no concrete conclusion can be drawn from these limited findings.

For instance, Ferreira (2006) extensively investigated the significance effect of monetary fundamentals on the exchange rates for Chile, South Korea, Malaysia, Mexico, South Africa, Thailand and Turkey from 1992 to 2002 using panel cointegration techniques. He considered the sticky price model to account for the price rigidities effect between developed and developing countries. The empirical evidence does not show any significant support to reject the hypothesis of no long run co-movement between exchange rates and monetary fundamentals across time and models.⁵ Therefore the finding casts doubt on the validity of the hypothesis introduced by McNown and Wallace (1994) who found significant co-movement between exchange rate and monetary fundamentals in some developing countries (Argentina, Chile and Israel). On the other hand, Wang and Wong (1997) used Kalman filtering techniques and ARCH models to address the issues of parameter instability and conditional variances to predict Japanese yen, Singapore dollar and Malaysian ringgit from 1973 to 1995. They found that the predictive power improves over 6 to 12 months forecasting horizons. The out-of-sample forecast errors are significantly lower compared to the naïve random walk model. Baharumshah and Masih (2005) further confirmed this finding using

⁴ Chinn (1998) stressed the importance of capital imperfect mobility and substitutability, and instability of money demand that are widespread in developing countries in monetary modelling in developing countries.

⁵ Panel cointegration techniques have been employed in order to mitigate the problem of small sample bias and to increase the power of the statistical test. However, Neely and Sarno (2002) cast doubt on the validity of across countries estimation since currency values in different countries may be driven by very different forces such as monetary policy and exchange rate regime.

cointegration techniques. They found substantial evidence of the strong predictive power of the monetary model, both for in-sample and out-of-sample forecast accuracy. Based on the standard root mean square error (RMSE) and the Theil's U statistics, their findings suggest that the structural model performs better than the random walk only when the current account is included into the VAR system. They also found the error-correction term in the exchange rate equation enters with a significantly negative coefficient. This could suggest that exchange rates converge to the equilibrium path over a longer period.

3.2 Monetary Models and Estimation Procedure

Theoretically, economists strongly believe that the exchange rate cannot deviate significantly from its 'fundamental value'. In other words, the exchange rate and the fundamental value are supposed to be cointegrated and one of the two variables will pull the other toward the equilibrium path. Therefore current deviations of the exchange rate from its fundamental value should help predict future exchange rate movements. As such, they may be represented in a typical dynamic error-correction framework:

$$\Delta s_{t+k} = s_{t+k} - s_t = \alpha_k + \lambda_k (s_t - f_t) + v_{t+k} \quad k = 1, 8, 12 \text{ and } 16 \quad (1)$$

where s_t is logarithm of the nominal domestic-currency price of one unit of foreign exchange at time t . f_t represents the fundamental value of the exchange rate. α_k is a constant and λ_k is the predictability parameter to be estimated. k is the forecast horizon (3 months or quarter of a year) and v_t is an *iid* disturbance term. If λ_k is smaller than 0, Equation 1 predicts that the exchange rate should depreciate when $s_t > f_t$ in order to revert toward the equilibrium path. A statistical test of predictability of exchange rate at horizon k is thus carried out based on the null hypothesis of no predictability, $H_0: \lambda_k = 0$ against the alternative hypothesis of predictability, $H_1: \lambda_k < 0$. There are at least two econometric procedures often used to estimate exchange rate predictability namely, traditional linear and non linear time series techniques. In this study we only considered the conventional linear time series methodology.

The estimation of Equation 1 was implemented in 2 steps. The first step consisted of obtaining the fundamental value f_t and the second was to estimate the forecasting regression. Specifically, first, we used Mark (1995) methodology to construct the fundamental value but with a few alterations to suite developing market characteristics. Instead of imposing theoretical value to the elasticity of money stock and income elasticity of money demand to $[1, -1]$ respectively, the fundamental value f_t will be constructed using the estimated elasticity of money stock and income elasticity of money demand from the estimated cointegrating coefficient of the Dynamic Ordinary Least Squares (DOLS) method. After constructing the fundamental values, then the forecasting estimation will be carried out employing the bootstrap procedure proposed by Mark (1995) and improved by Kilian (1999) under a constrained error-correction specification.

3.2.1 Construction of the Fundamental Values

The fundamental values f_t was constructed using cointegrating coefficients estimated by DOLS regression using the following specification:⁶

$$s_t = \alpha + \beta f_t + \sum_{j=-q}^q \delta_j \Delta f_{t-j} + \varepsilon_t \quad (2)$$

where f_t is a vector of fundamental variables obtained from either one of the following three monetary models:

Flexible price model

$$f_t = [(m_t - m_t^*), (y_t - y_t^*)] \quad (3)$$

Sticky price model

$$f_t = [(m_t - m_t^*), (y_t - y_t^*), (i_t - i_t^*), (\pi_t - \pi_t^*)] \quad (4)$$

Relative price model

$$f_t = [(m_t - m_t^*), (y_t - y_t^*), (i_t - i_t^*), (\pi_t - \pi_t^*), (p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*})] \quad (5)$$

where m , y , i , π and p in Equations 3, 4 and 5 represent the logarithm of money stock, the logarithm of real income, nominal interest rate, the CPI inflation rate and overall prices which include T , tradable, and N , non tradable goods, respectively. An asterisk indicates foreign markets. β in Equation 2 is a vector of parameters of the corresponding monetary models (flexible price, $[\beta_m, \beta_y]$; sticky price, $[\beta_m, \beta_y, \beta_i, \beta_\pi]$; and relative price, $[\beta_m, \beta_y, \beta_i, \beta_\pi, \beta_p]$). The β_m represents the elasticity of money stock, β_y is the income elasticity of money demand, β_p is the relative price elasticity, β_i and β_π are the interest and inflation semi-elasticity, respectively. The anticipated sign for the estimated coefficients are β_m, β_p and $\beta_\pi > 0$, while β_y and $\beta_i < 0$. Δ is difference operator and following Stock and Watson (1993), we set the number of leads and lags of the regressor (q) in the DOLS estimator of Equation 2 equal to three ($q = 3$). We used Newey-West procedure to compute robust standard errors.

⁶ For $I(1)$ series with one cointegration relation, the DOLS estimation procedure produces efficient estimates of the cointegrating vector.

The estimated cointegrating coefficients, $\hat{\beta}$ s in Equation 2, were then used to construct the fundamental values based on the following models:

Flexible price

$$\hat{f}_t = \hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) \quad (6)$$

Sticky price

$$\hat{f}_t = \hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) - \hat{\beta}_i(i_t - i_t^*) + \hat{\beta}_\pi(\pi_t - \pi_t^*) \quad (7)$$

Relative price

$$\begin{aligned} \hat{f}_t = & \hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) - \hat{\beta}_i(i_t - i_t^*) + \hat{\beta}_\pi(\pi_t - \pi_t^*) \\ & + \hat{\beta}_p[(p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*})] \end{aligned} \quad (8)$$

Deriving fundamental values using the standard flexible price monetary model (Equation 6) is the most common procedure that has been extensively used by most of the researchers in the area, Mark (1995) and Kilian (1999) among others.⁷ However, it is less appropriate in the case of developing countries since it requires domestic and foreign asset to be perfect substitutes and uncovered interest parity (UIP) condition to hold in the markets.

In this paper, we also considered two extensions of the basic monetary model as suggested by Chinn (1998). First, following the work of Dornbusch (1976) and Frankel (1979) we considered a monetary model that incorporates short-term price rigidities (Equation 7). This model incorporates variables that allow for short-run price stickiness that violates the Purchasing Power Parity (PPP) hypothesis. In addition, the relationship includes interest rates in order to capture the short-term liquidity effect of the monetary policy. Second, we considered relative price movements by including the tradable and non tradable goods within and across countries. Following Balassa (1964) and Samuelson (1964), the relative prices model is driven by relative differentials in productivity in the tradable and non tradable sectors as presented in Equation 8. These two approaches, i.e. the sticky price monetary model and the relative price Balassa-Samuelson model, are expected to represent better the fundamental values of developing economies. Furthermore, the inclusion of sticky prices and the Balassa-Samuelson effect in \hat{f}_t could be crucial for finding cointegration evidence in developing countries.

Equation 1, combined with the structural models discussed above, resulted in the following predictability equations:

Model 1:

$$\Delta s_{t+k} = \alpha + \lambda_k [s_t - (\hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*))] + \varepsilon_{t+k} \quad (9)$$

⁷Mark (1995) and Kilian (1999) imposed value of [1, -1] to β_m and β_y respectively.

Model 2:

$$\Delta s_{t+k} = \alpha + \lambda_k [s_t - (\hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) - \hat{\beta}_i(i_t - i_t^*) + \hat{\beta}_\pi(\pi_t - \pi_t^*))] + \varepsilon_{t+k} \quad (10)$$

Model 3:

$$\Delta s_{t+k} = \alpha + \lambda_k [s_t - (\hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) - \hat{\beta}_i(i_t - i_t^*) + \hat{\beta}_\pi(\pi_t - \pi_t^*) + \hat{\beta}_p((p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*})))] + \varepsilon_{t+k} \quad (11)$$

3.2.2 Forecasting Regression

We considered in-sample and out-of-sample forecasts to evaluate the accuracy of the monetary model in predicting exchange rate movements. Analysis of the in-sample forecast (based on full sample from 1984Q1 to 2005Q4) of the monetary models (Models 1, 2 and 3) has been compared to random walk model of Equation 12 (as a benchmark)

$$s_{t+k} - s_t = d_k + \varepsilon_{t+k} \quad k = 1, 8, 12 \text{ and } 16 \quad (12)$$

of the corresponding k and tested for $H_0 : \lambda_k = 0$ against $H_1 : \lambda_k < 0$ or based on joint test of all forecast horizons as $H_0 : \lambda_k = 0 \forall k$ against $H_1 : \lambda_k < 0$ for some k . On the other hand, for out-of-sample forecast, we used prediction mean-squared error of Equations 9, 10, 11 and 12 from the sequence of recursive forecasts to evaluate the Theil's U -statistic and DM statistic of Diebold and Mariano (1995) with and without drift. Specifically, the estimation started from 1984Q1 to 1995Q4. To generate the next forecast k , the estimation sample was updated by one period 1996Q1 for $k = 1$, 1997Q4 for $k = 8$, 1998Q4 for $k = 12$ and 1999Q4 for $k = 16$. The procedure was repeated until we reached the end of the sample in 2005Q4.

However, the forecasting exercise based on Models 1, 2 and 3 involved econometric difficulties. First, the error-correction representation was only appropriate under the assumption of stationarity of the error correction term $(s_t - f_t)$. This is because the asymptotic null distribution of test statistics for λ_k depends on whether the error-correction term is stationary or not, as discussed in Cavanagh *et al* (1995) and Valkanov (2003).

Another econometric problem is that forecasting involves future horizons k ; when $k > 1$, the dependent variable $(s_{t+k} - s_t)$ represents overlapping sums of the original series that may result in high persistency of the error correction term. In this case, statistical inference should be handled with care since the in-sample R^2 and the t -statistics do not converge to a well-defined asymptotic distribution and the estimated coefficient, $\hat{\lambda}_k$, is biased away from zero due to size distortions. This bias is in favour of finding predictability

as the forecast horizon (k) increases (see among others Mark and Sul 2001; Berkowitz and Giorgianni 2001), for detailed discussions on the subject matter).

To mitigate the above discussed problems we considered the bootstrap technique proposed by Kilian (1999) to approximate the finite sample distribution of the test statistic under the null hypothesis of no exchange rate predictability. This approach consisted, first, of estimating the Data-Generating Process (DGP) under the null of no predictability for the Constrained Vector Error Correction Model (VECM)

$$\Delta s_t = \alpha_s + u_{1,t} \quad (13)$$

and

$$\Delta f_t = \alpha_f - h_2(f_{t-1} - s_{t-1}) + \sum_{j=1}^{q-1} \xi_j^{21} \Delta s_{t-j} + \sum_{j=1}^{q-1} \xi_j^{22} \Delta f_{t-j} + u_{2,t} \quad (14)$$

using constrained Estimated Generalised Least Squares (EGLS) technique with all coefficients but α_s set equal to zero. The system also required the restriction of $h_2 < 0$ to be satisfied to ensure estimation stability. The lag order q was determined under H_0 using AIC criterion.⁸

Second, after estimating Equations 13 and 14, a sequence of $\{s_t^*, f_t^*\}$, pseudo observations could be generated under the assumption of *i.i.d* innovations using cumulative sums of the realisations of the bootstrap data-generating process. The process was initialised by specifying $(f_{t-1}^* - s_{t-1}^*) = 0$ and $\Delta s_{t-j}^* = 0$ and $\Delta f_{t-j}^* = 0$ for $j = q-1, \dots, 1$ and discarding the first 500 observations. The pseudo innovation term $u_t^* = (u_{1t}^*, u_{2t}^*)'$ was random and drawn with replacement from the set of observed residuals $\hat{u}_t = (\hat{u}_{1t}, \hat{u}_{2t})'$. The process was repeated 2000 times. Third, these $\{s_t^*, f_t^*\}$, of 2000 bootstrap replication were used to estimate the following long-horizon regression:

$$s_{t+k}^* - s_t^* = \alpha_k^* + \lambda_k^* (s_t^* - f_t^*) + v_{t+k}^* \quad k = 1, 4, 8, 12, 16 \quad (15)$$

Finally, the empirical distribution of these 2000 replications of the bootstrap test statistics were used to determine the p -value of the $t(20)$, $t(A)$, U , $DM(20)$, and $DM(A)$ of Equations 9, 10 and 11.

With regard to the potential problem of the serial correlation of the error term due to $k > 1$, we adopted two approaches. First we used Newey-West corrected t -statistics by setting the truncation lags to 20 since the longest forecast horizon was 16. Second, we used a data-dependent formula provided by Andrews (1991) under a univariate AR(1) as an approximating model. As a result, the statistical inference was robust to highly persistent or

⁸ For further detailed explanation on the estimation procedures, please refer to appendix in Kilian (1999).

Table 1. Economic fundamentals for selected developing countries from 1984q1 to 2005q4

Country	Exchange rate regime		Income volatility	Inflation	Total trade (% GDP)
	Lowest	Highest			
Chile	1	2	2.75	11.63	60.32
Israel	1	2	2.27	41.12	83.41
Morocco	1	1	4.84	4.11	60.84
Philippines	1	3♣	3.89	9.91	79.11
South Africa	1	3♦	2.59	9.99	48.50
Thailand	1	2	4.78	3.61	90.73
Tunisia	1	2	2.60	5.00	87.36
Uruguay	1	2	5.08	43.41	42.22
United States	1	1	1.53	3.11	21.64

Classification of exchange rate regime is based on Levy-Yeyati and Sturzenegger (2003; 2005). The index ranges from 1 = float; 2 = intermediate; and 3 = fixed. We included Philippines and S. Africa since they had a fixed regime only for these three years ♣ 1987, 1993 and 1996 and ♦ 1990, 1993 and 1995, respectively. The indices for the remaining countries and years are either 1 or 2. Income volatility is the standard deviation of the growth rate of GDP per capita. Inflation is a measure of mean inflation over the sample period. Total trade is an average of total import and export per GDP.

had near-spurious regression problems because it had the ability to automatically adjust the critical values to the increase in dispersion of the finite sample distribution of the test statistic for different lag structures and estimation procedures.

3.3 Data

In the present case, which is limited by the availability of fully liberalised developing economies, we constrained ourselves to markets that satisfied the assumptions of the monetary model i.e. floating exchange rate regime and relatively open capital markets over a long period. Based on Levy-Yeyati and Sturzenegger (2003; 2005), and supplemented by ratios of total external trade to GDP (see Table 1), we chose the following 8 developing economies: Chile, Israel, Morocco, Philippines, South Africa, Thailand, Tunisia and Uruguay (along with the US economy as a base market).⁹ Levy-Yeyati and Sturzenegger (2003) classified 3 *de-facto* exchange rate regimes: float, intermediate and fixed. We chose only markets that were under float or intermediate regimes for the whole sample periods.¹⁰ Float

⁹ The definition of developing market is based on the International Financial Cooperation (IFC). For more detailed explanation refer to Global Economic Prospects and the Developing Countries, World Bank (2002), among others.

¹⁰ We included Philippines and South Africa into our sample though these two economies had fixed exchange rate regime for the following years: 1987, 1993, 1996 and 1990, 1993, 1995 respectively. Countries that experienced more than three years of fixed exchange rates regime were excluded from the analysis. Full version of exchange rate regime classification can be accessed from <http://200.32.4.58/~ely/index.html>.

and intermediate regimes also indirectly indicate that the markets are not only open but characterised by little market frictions and government intervention. As defined by Levy-Yeyati and Sturzenegger (2003), float and intermediate regimes are characterised by indices of low reserve volatility together with high exchange rate volatility. Low volatility of reserves is considered an indicator of less government intervention in the monetary policy. Therefore countries that have adopted a hard peg exchange regime, like China and Malaysia, or excessive capital control, like Korea, were excluded from the analysis.

The variables considered in our monetary model were end-of-period quarterly nominal exchange rates expressed as the US dollar per developing country currency to proxy the nominal exchange rate (s_t), the money stock M2 to measure money supply (m_t), the Gross Domestic Product (GDP) to proxy real output (y_t), the Consumer Price Index (CPI) as broad deflator (π_t), short term interest rate is proxied by inter-bank deposit interest rates (i_t), and the relative price of tradable and non tradable price deflator (p_t) is proxied by the ratio of CPI and Producers Prices Index (PPI) or Wholesale Price Index (WPI). The sample period considered in the analysis was from 1984Q1 to 2005Q4 and retrieved from either Datastream® or the IMF's International Financial Statistics. All variables except interest rates were converted to natural logarithms.

4. Results

Unlike earlier studies (for instance, Mark 1995; Kilian 1999), this paper does not impose a theoretical value for the cointegrating coefficients in constructing the fundamental values (\hat{y}_t). Instead, we used the estimated parameters obtained from DOLS regressions of Equation 2. Table 2 shows the estimated cointegrating coefficients that were used in constructing the fundamental values for all models and markets.

We computed the Theil's U -statistics (the ratio of RMSE from two competing models—monetary versus random walk), the t -statistics and the Diebold-Mariano, (DM) statistics to assess the performance of exchange rate forecast using Models 1, 2, and 3.¹¹ The estimation results are presented in Table 3a and 3b for the drift-less random walk benchmark model and in Table 3c and 3d for the random walk with a drift term. All the test results are presented in the form of bootstrap p -values based on 2000 replications. We were particularly interested in testing (in-sample) the hypothesis that $\lambda_k < 0$, and the out-of-sample performance based on one-step ahead the Diebold-Mariano DM test statistics and Theil's U -statistics. Long horizon predictability arises if the p values indicate increasing significance as the horizon k becomes larger. We were also interested in testing the joint significance of $\lambda_k = 0$ for all k at 10 per cent level.

Based on these criteria, the results show that only two countries (Israel and Uruguay) provide strong support for long horizon out-of-sample predictability. For Israel, the forecast accuracy improves for longer horizons. This is evident from the U -statistics that are significant at $k = 12$ and 16 under the no drift sticky price model. In addition, the p value of

¹¹ The estimation procedures were conducted using the MATLAB code provided by Lutz Kilian which is available in the *Journal of Applied Econometrics* data and code archive 1999 Volume 5.

Table 3a. Results of the VEC bootstrap model with no drift

Country	Horizon	Flexible Price Model				Sticky Price Model				Relative Price Model						
		$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$
Chile	1	0.059	0.017	0.071	0.081	0.074	0.029	0.010	0.149	0.133	0.140	0.043	0.013	0.064	0.076	0.067
	8	0.157	0.073	0.295	0.266	0.308	0.071	0.027	0.230	0.204	0.225	0.123	0.055	0.222	0.207	0.237
	12	0.198	0.137	0.251	0.283	0.282	0.080	0.046	0.178	0.178	0.178	0.149	0.111	0.196	0.197	0.196
	16	0.241	0.216	0.287	0.625	0.389	0.081	0.071	0.174	0.135	0.138	0.146	0.128	0.207	0.229	0.218
	Joint	0.285	0.190	0.319	0.296	0.293	0.120	0.077	0.248	0.210	0.211	0.200	0.153	0.294	0.266	0.261
Uruguay	1	0.353	0.165	0.990	0.542	0.426	0.353	0.163	0.990	0.539	0.419	0.373	0.183	0.970	0.538	0.416
	8	0.340	0.330	0.980	0.668	0.720	0.336	0.327	0.990	0.663	0.722	0.359	0.349	0.990	0.660	0.724
	12	0.285	0.301	0.980	0.639	0.674	0.290	0.303	0.980	0.648	0.688	0.319	0.340	0.980	0.658	0.689
	16	0.210	0.221	0.980	0.638	0.675	0.216	0.232	0.980	0.639	0.673	0.239	0.267	0.980	0.631	0.660
	Joint	0.283	0.300	0.980	0.732	0.612	0.288	0.300	0.990	0.738	0.616	0.310	0.330	0.980	0.731	0.618
Philippines	1	0.577	0.613	0.487	0.567	0.423	0.750	0.755	0.384	0.396	0.316	0.661	0.670	0.391	0.449	0.333
	8	0.789	0.787	0.145	0.124	0.134	0.805	0.807	0.212	0.226	0.226	0.795	0.793	0.205	0.216	0.212
	12	0.811	0.812	0.208	0.233	0.220	0.838	0.839	0.325	0.544	0.448	0.816	0.813	0.281	0.474	0.386
	16	0.883	0.887	0.211	0.183	0.193	0.874	0.879	0.360	0.545	0.593	0.867	0.868	0.320	0.482	0.472
	Joint	0.773	0.809	0.289	0.266	0.278	0.851	0.852	0.475	0.424	0.417	0.814	0.819	0.513	0.410	0.406
Thailand	1	0.641	0.500	0.988	0.602	0.506	0.636	0.581	0.551	0.735	0.628	0.595	0.500	0.632	0.689	0.649
	8	0.748	0.751	0.265	0.676	0.550	0.722	0.720	0.327	0.765	0.606	0.752	0.747	0.456	0.968	0.896
	12	0.795	0.797	0.325	0.888	0.776	0.812	0.812	0.616	0.618	0.658	0.803	0.802	0.656	0.638	0.664
	16	0.829	0.827	0.460	0.636	0.673	0.842	0.844	0.654	0.529	0.573	0.852	0.852	0.704	0.517	0.554
	Joint	0.843	0.837	0.684	0.795	0.688	0.813	0.801	0.623	0.863	0.821	0.801	0.794	0.698	0.819	0.828

Note: The figures under $t(20)$, $t(A)$, U , $DM(20)$ and $DM(A)$ headings are bootstrap p -values for the VEC model with or without drift (Kilian 1999). Flexible price model, sticky price model and Balassa-Samuelson effect model have been considered to construct the fundamental variables. $t(20)$ refers to t -statistic for the slope coefficient in the long-horizon regression with robust standard errors calculated based on a fixed truncation lag of 20. $t(A)$ refers to the case of standard errors using Andrew (1991) rule. DM and U refer to the corresponding Diebold-Mariano and Theil's U -statistics (ratio of out-of-sample and random walk model) respectively. Results are shown for alternative forecast horizons $k = 1$ -, 8-, 12- and 16-quarter. Joint refers to the p -value for the joint test statistics for all horizons. Boldface p values denote significance at the 10 per cent level.

Table 3b. Results of the VEC bootstrap model with no-drift

	Country	Horizon	Flexible price model				Sticky price model				Relative price model					
			$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$
Israel	1	0.547	0.493	0.134	0.149	0.150	0.146	0.122	0.177	0.129	0.126	0.424	0.359	0.102	0.121	0.124
	8	0.480	0.464	0.184	0.199	0.198	0.135	0.126	0.204	0.174	0.173	0.391	0.369	0.230	0.204	0.204
	12	0.321	0.319	0.051	0.122	0.124	0.066	0.066	0.047	0.086	0.090	0.256	0.246	0.064	0.123	0.124
	16	0.242	0.284	0.081	0.133	0.141	0.060	0.073	0.085	0.104	0.114	0.186	0.225	0.080	0.109	0.116
	Joint	0.313	0.358	0.097	0.205	0.211	0.086	0.092	0.086	0.151	0.152	0.245	0.279	0.094	0.167	0.177
Morocco	1	0.142	0.168	0.336	0.357	0.360	0.142	0.187	0.311	0.402	0.353	0.058	0.058	0.096	0.094	0.086
	8	0.072	0.073	0.146	0.143	0.145	0.155	0.151	0.261	0.819	0.806	0.104	0.120	0.214	0.240	0.230
	12	0.183	0.186	0.195	0.166	0.165	0.262	0.248	0.210	0.558	0.684	0.181	0.185	0.205	0.180	0.180
	16	0.197	0.203	0.313	0.376	0.407	0.284	0.268	0.332	0.411	0.508	0.176	0.185	0.335	0.380	0.422
	Joint	0.190	0.196	0.275	0.291	0.287	0.294	0.307	0.453	0.663	0.558	0.257	0.248	0.293	0.291	0.282
S. Africa	1	0.607	0.594	0.460	0.188	0.202	0.724	0.722	0.382	0.190	0.197	0.611	0.612	0.414	0.183	0.192
	8	0.496	0.467	0.536	0.271	0.290	0.583	0.556	0.414	0.255	0.273	0.489	0.464	0.506	0.255	0.278
	12	0.569	0.544	0.355	0.238	0.250	0.651	0.637	0.266	0.218	0.224	0.566	0.538	0.340	0.240	0.245
	16	0.689	0.683	0.212	0.215	0.215	0.751	0.749	0.134	0.175	0.170	0.679	0.675	0.188	0.204	0.202
	Joint	0.676	0.644	0.329	0.344	0.343	0.750	0.725	0.169	0.279	0.263	0.662	0.634	0.259	0.320	0.313
Tunisia	1	0.525	0.581	0.780	0.939	0.767	0.187	0.239	0.121	0.121	0.121	0.415	0.451	0.608	0.326	0.355
	8	0.508	0.519	0.659	0.521	0.566	0.268	0.260	0.316	0.231	0.236	0.305	0.325	0.696	0.564	0.626
	12	0.567	0.538	0.586	0.592	0.633	0.338	0.330	0.192	0.189	0.190	0.398	0.360	0.703	0.887	0.900
	16	0.705	0.703	0.295	0.295	0.306	0.497	0.497	0.095	0.135	0.121	0.587	0.591	0.331	0.313	0.328
	Joint	0.538	0.626	0.815	0.474	0.510	0.395	0.401	0.119	0.196	0.171	0.382	0.438	0.671	0.538	0.565

Note: Refer to note in Table 3a

Table 3c. Results of the VEC bootstrap model with drift

Country	Horizon	Flexible price model				Sticky price model				Relative price model						
		$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$
Chile	1	0.030	0.010	0.024	0.043	0.026	0.044	0.013	0.013	0.040	0.024	0.054	0.016	0.012	0.059	0.037
	8	0.070	0.026	0.237	0.221	0.222	0.122	0.051	0.252	0.225	0.228	0.149	0.067	0.349	0.266	0.271
	12	0.080	0.047	0.582	0.373	0.433	0.147	0.111	0.596	0.383	0.408	0.200	0.136	0.694	0.422	0.454
	16	0.080	0.069	0.771	0.484	0.495	0.144	0.126	0.798	0.515	0.507	0.237	0.207	0.859	0.572	0.555
	Joint	0.120	0.077	0.259	0.281	0.259	0.197	0.151	0.246	0.262	0.233	0.280	0.193	0.281	0.320	0.289
Uruguay	1	0.098	0.029	0.003	0.025	0.018	0.099	0.027	0.003	0.025	0.017	0.102	0.028	0.003	0.027	0.017
	8	0.069	0.056	0.126	0.120	0.121	0.068	0.057	0.126	0.119	0.122	0.071	0.058	0.125	0.120	0.122
	12	0.056	0.044	0.124	0.122	0.123	0.055	0.045	0.120	0.118	0.121	0.057	0.050	0.121	0.120	0.121
	16	0.043	0.039	0.153	0.150	0.150	0.043	0.039	0.152	0.151	0.153	0.047	0.040	0.151	0.148	0.151
	Joint	0.045	0.040	0.121	0.102	0.098	0.045	0.040	0.120	0.099	0.095	0.049	0.041	0.122	0.100	0.095
Philippines1	1	0.746	0.752	0.778	0.843	0.789	0.628	0.650	0.700	0.588	0.744	0.310	0.344	0.230	0.412	0.583
	8	0.803	0.803	0.721	0.772	0.759	0.778	0.774	0.464	0.654	0.599	0.620	0.617	0.171	0.912	0.743
	12	0.833	0.836	0.813	0.910	0.920	0.833	0.833	0.592	0.879	0.918	0.602	0.600	0.197	0.922	0.930
	16	0.870	0.877	0.874	0.853	0.896	0.877	0.879	0.631	0.828	0.864	0.539	0.542	0.244	0.905	0.854
	Joint	0.851	0.849	0.834	0.942	0.903	0.739	0.764	0.726	0.772	0.735	0.466	0.531	0.320	0.633	0.757
Thailand 1	1	0.636	0.581	0.593	0.711	0.720	0.720	0.584	0.942	0.712	0.624	0.585	0.445	0.993	0.675	0.586
	8	0.722	0.719	0.390	0.898	0.867	0.798	0.804	0.429	0.948	0.877	0.677	0.680	0.291	0.259	0.272
	12	0.811	0.811	0.719	0.685	0.728	0.858	0.858	0.813	0.781	0.769	0.725	0.727	0.354	0.822	0.551
	16	0.843	0.843	0.792	0.626	0.670	0.891	0.891	0.919	0.644	0.671	0.754	0.759	0.511	0.645	0.667
	Joint	0.813	0.801	0.721	0.911	0.903	0.865	0.850	0.917	0.848	0.757	0.779	0.768	0.485	0.457	0.471

Note: Refer to note in Table 3.

Note: Refer to note in Table 3a

Table 3d. Results of the VEC bootstrap model with drift

Country	Horizon	Flexible Price Model				Sticky Price Model				Relative Price Model			
		$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	U
		$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	$t(A)$	U	$DM(20)$	$DM(A)$	$t(20)$	U
Israel	1	0.148	0.114	0.435	0.267	0.455	0.398	0.313	0.217	0.217	0.224	0.532	0.476
	8	0.140	0.128	0.502	0.284	0.302	0.363	0.335	0.441	0.320	0.329	0.464	0.450
	12	0.068	0.069	0.211	0.223	0.223	0.228	0.224	0.282	0.277	0.277	0.302	0.306
	16	0.063	0.070	0.389	0.284	0.288	0.170	0.203	0.293	0.301	0.301	0.234	0.272
	Joint	0.089	0.090	0.383	0.427	0.426	0.220	0.250	0.413	0.438	0.438	0.300	0.331
Morocco	1	0.142	0.189	0.317	0.317	0.282	0.058	0.058	0.060	0.075	0.061	0.142	0.169
	8	0.153	0.151	0.236	0.232	0.234	0.104	0.104	0.135	0.177	0.185	0.072	0.073
	12	0.261	0.247	0.207	0.198	0.194	0.181	0.185	0.143	0.201	0.185	0.183	0.186
	16	0.282	0.266	0.497	0.391	0.424	0.175	0.184	0.543	0.401	0.435	0.197	0.203
	Joint	0.295	0.307	0.340	0.358	0.347	0.256	0.247	0.251	0.338	0.331	0.190	0.196
S. Africa	1	0.723	0.721	0.676	0.481	0.416	0.586	0.542	0.490	0.403	0.325	0.586	0.542
	8	0.584	0.555	0.655	0.501	0.532	0.543	0.522	0.529	0.376	0.395	0.543	0.522
	12	0.651	0.635	0.664	0.455	0.495	0.615	0.591	0.429	0.358	0.372	0.615	0.591
	16	0.750	0.750	0.684	0.545	0.589	0.747	0.747	0.414	0.356	0.367	0.747	0.747
	Joint	0.748	0.722	0.782	0.748	0.674	0.665	0.630	0.528	0.546	0.441	0.665	0.630
Tunisia	1	0.186	0.240	0.134	0.130	0.129	0.415	0.451	0.796	0.771	0.671	0.524	0.579
	8	0.265	0.261	0.562	0.618	0.693	0.305	0.327	0.852	0.968	0.989	0.506	0.520
	12	0.339	0.330	0.484	0.423	0.461	0.397	0.362	0.908	0.994	0.999	0.565	0.540
	16	0.500	0.500	0.382	0.359	0.372	0.588	0.592	0.775	0.973	0.987	0.706	0.704
	Joint	0.397	0.401	0.469	0.457	0.455	0.381	0.439	0.863	0.960	0.901	0.539	0.627

Note: Refer to note in Table 3a

the joint test of the *Theil's U*-statistics is also significant. However none of the test statistics for Israel are significant when a drift term is considered in the models. In the case of Uruguay, the monetary models with a drift predict better the exchange rate movements. The joint test of *DM(20)* statistics for sticky price model and *DM(A)* for all three models with a drift are significant compared to none for the driftless case.

The result shows that there is evidence of the short horizon ($k = 1$ and 8) predictability of Chile, Uruguay and Morocco under the monetary models with a drift term. The out-of-sample test statistics (for $k = 1$) of all models are significant for Chile and Uruguay but only sticky price model fits the Moroccan market. Another obvious finding from the analysis is that the Chilean, Israeli and Uruguayan markets also provide significant support for in-sample predictability. The p values of $t(A)$ and $t(20)$ for some of the λ_k are significant (in the case of the Uruguayan market and the in-sample predictability test statistics are significant for all models with drift term). For the remaining countries (Philippines, Thailand, South Africa and Tunisia), no predictability has been detected in the analysis.

A number of interesting observations can be drawn from the results discussed above. First, the two countries (Israel and Uruguay) for which we find support of long-horizon predictability are characterised by high inflation (see for instance Braumann 2000) for high inflation countries classification and Table 1 for comparison between markets under study). The results confirm the earlier proposition made by McNown and Wallace (1994) and Rogoff (1999a) who argue that forecast accuracy using monetary models should be higher in countries with unstable macroeconomic fundamentals.

Second, inclusion of a drift term in the estimation has eliminated predictability from the Israeli market. The opposite holds for Uruguay where predictability arises when the benchmark is the random walk with drift. This shows the importance of considering drift or no drift in the estimation, as argued by Kilian (1999). Third, considering alternative monetary models (sticky price and relative price models) has proved to be useful in the process of predicting exchange rates movements in developing countries. At least the sticky price model appears to be superior to the standard flexible price and the relative price Balassa-Samuelson model. This finding is similar to that of Chinn (1998) who suggested the superiority of the sticky price model over the relative price Balassa-Samuelson for Philippines peso and Thailand baht.

Finally, the finding of short-term predictability ($k = 1$ and 8) for Chile, Uruguay and Morocco is relatively surprising. This could be presumably a result of the instantaneous exchange market reaction to the instability of economic fundamentals. The evidence is in favour with the growing literature on the integration of the currency market (Francis *et al.* (2002) and equity market (Frankel *et al.* 2004; Golstein *et al.* 2000) in developing economies.¹²

¹² Investors could be very prudent on the issue of economic and non economic uncertainty in developing countries. Studies have shown that the risk of international loan default is very high (Catao and Sutton, 2002, among others). Consequently, for highly unstable developing countries, they are unlikely to receive huge amounts of long-term debt. For instance, Velev (2007) has shown that most of the unstable developing countries have received significant higher, short-term maturity debt compared to the relatively stable developing economies. Uruguay which is more unstable has received 63 per cent short-term credit from the US banks compared to only 20 per cent for Peru. Instability has a significant role in determining short- or long-term credit to the developing countries.

The linkages between markets are further speeded up by the rapid development in information technology.¹³

5. Conclusion

We considered developing countries that are open and adopted floating exchange rate regimes to investigate the exchange rate forecastability puzzle using monetary models. The motivation for this study is based on the hypothesis proposed by McNown and Wallace (1994) and Rogoff (1999a). The hypothesis states that exchange rate predictability should be better off in countries with unstable monetary fundamentals. In addition to the standard flexible price model, we considered two alternatives approaches that account for sticky and relative prices. The method of Kilian (1999) was employed to reduce problems in the long horizon finite sample forecasting estimations.

Based on Levy-Yeyati and Sturzenegger (2003; 2005), eight developing countries were chosen for the analysis to gain insights into exchange rate forecastability. The results suggest that the inclusion of fundamental values derived from the sticky price monetary model appear to improve the out-of-sample forecast accuracy of the exchange rate determination models for four developing economies of Chile, Israel, Morocco and Uruguay. Empirical evidence is in favour of the hypothesis that markets with unstable monetary fundamentals such as high inflation have higher forecast accuracy compared to the ransom walk model.

Overall, predictability of exchange rates in developing countries is very sensitive to the selection of appropriate models and the results are country specific in nature. For future research on developing countries on the same issue, it may be fruitful to explore the potential of short- or long-term forecast accuracy using non linear specifications.

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¹³ The plausibly true channels of the linkage are trade linkages (Glick and Rose 1999), 'common lender' or stock market (Kaminsky and Reinhart, 2001 and Caramazza *et al.*, 2000), and 'common macro-economic weaknesses' (Eichengreen *et al.* 1996).

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