Exchange Rates And Fundamentals: A Non-Linear Relationship?

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Abstract

We test whether the relationship between changes in the nominal exchange rate and changes in its underlying fundamentals has non-linear features. In order to do so, we extend the Markov switching model as proposed by McConnell and Perez Quiros (2000) and Dewachter (2001) and test it using a sample of low and high inflation countries. The empirical analysis shows that for the high inflation countries the relationship between news in the fundamentals and the exchange rate changes is stable and significant. This is not the case, however, for the low inflation countries, where frequent regime switches occur. We develop a non-linear model based on the existence of transactions costs that could explain our empirical findings. We find that this simple non-linear model is capable of replicating the empirical evidence uncovered in this paper.

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1 Introduction

Exchange rate economics has gone through different stages. The early theoretical models were developed mainly in the 1970s (monetary model, Dornbusch model, portfolio balance model, and others). These 'first generation' models led to testable propositions in which the changes in the exchange rate are linearly related to news in the fundamentals (money stocks, prices, output, current accounts, etc.). After intensive empirical testing it is fair to conclude that the first generation models were soundly rejected by the data, at least for the exchange rates of countries experiencing relatively low levels of inflation. Three serious anomalies of the first generation models were detected.

First, in their celebrated empirical studies Meese and Rogoff (1983), (1988) found that the random walk forecast typically outperforms a forecast based on the first generation models.[1] Although occasionally some researchers have claimed that their model could beat the random walk, the scientific consensus today is that the Meese and Rogoff results still stand at least as far as short-term (one-period ahead) forecasting is concerned. There is some recent evidence, however, indicating that non-linear models are capable of beating the random walk at medium term horizons (see Kilian and Taylor, 2003).

A second anomaly detected in the empirical literature is the following. Since the start of the floating exchange rate regime the variability of the exchange rates (both nominal and real) has increased dramatically. At the same time there is no evidence to be found that the variability of the fundamentals identified by the theoretical models has increased compared to the fixed exchange rate period (see Baxter and Stockman, 1989 and Flood and Rose, 1995). This is in contradiction with the first generation models, which imply that the variability of the exchange rate can only increase when the variability of the underlying fundamental variables increases. This result has led to the view that the variability of the exchange rates is largely disconnected from the variability of the underlying fundamentals. In their recent paper Obstfeld and Rogoff (2000) have identified this phenomenon to be one of the six major puzzles in international macroeconomics.

A third empirical anomaly relates to the 'news' aspect of the first generation models. The rational expectations assumption underlying the first generation models implies that the exchange rates can only change at any given moment of time as a result of 'news' in the fundamentals. Empirical analysis using structural VARs however comes to conflicting answers on this. On the one hand Clarida and Gali (1994) and Farrant and Peersman (2005) find that a substantial part of real exchange rate fluctuations can in fact be explained by its underlying fundamentals. At the same time, De Boeck (2000) and Altavilla (2000) find that unanticipated shocks in the fundamental variables explain only a small fraction of the unanticipated changes in the exchange rates. Typically over forecast horizons of up to one year, they find that news in output, inflation, and interest rates explains less than 5% of the total unanticipated variance of the exchange rate. About 95% of the latter is attributable to the news in the exchange rate itself.[2] From

[1] There is some evidence that when forecasting over a longer horizon, say, more than one year, fundamentals based models sometimes outperform the random walk (see e.g. Mark, 1995).

[2] Again there is some evidence that over longer forecast horizons, the news in fundamentals becomes more important. It remains relatively low, however, remaining far below explaining 50% of the total variance.

this evidence it appears that the first generation models in which the exchange rate is driven by news in the fundamentals in a linear way must be called into question as a representation of the foreign exchange market.

The rejection of the first generation models of the exchange rate has led researchers into different directions. The first one has led to what one could call the 'second generation' models, as exemplified by Obstfeld and Rogoff (1996). In these models the starting point is utility maximisation of a representative agent. These models typically lead to the conclusion that the coefficients of the reduced form equations of the first generation models do not have to be constant. These coefficients vary as a result of the underlying stochastic disturbances and of changing policy regimes.

This is an important insight. The trouble, however, is that the 'second generation' models have led to few testable propositions that would allow for their refutation. As long as these testable propositions are not formulated it is difficult to evaluate the scientific strength of these 'second generation' models.

A second direction taken by researchers in their search for an alternative to the 'first generation' models has been to introduce nonlinearities into the model (see De Grauwe and Dewachter, 1993, Frankel and Froot, 1990, Kilian and Taylor, 2001, and Kurz and Motolese, 2001). These models are characterised by the existence of several agents using different information sets (e.g. chartists and fundamentalists) and/or by the existence of transactions costs. The insight provided by these models is that they predict frequent structural breaks in linear exchange rate equations, and that they allow for changes in the exchange rates that are unrelated to news about the underlying fundamentals.

In this paper we analyse the (possibly non-linear) nature of the relationship between exchange rate changes and the changes in the underlying fundamentals. More specifically we test whether this relationship is subject to regime switches over time. In order to do so, we use a version of the Markov-switching autoregressive model popularised by Hamilton (1989). We perform this analysis both on data of low inflation and high inflation countries. This comparison between low and high inflation countries will allow us to gain additional insight about the nature of the relation between exchange rates and the fundamentals. Several recent studies have used similar Markov-switching models or ESTAR models to detect non-linearities in the exchange rate dynamics (see e.g. Peel and Taylor (2000) and Taylor, Peel and Sarno (2001)). The value added of our study is that it makes a distinction between low and high inflation countries. As will be shown, the level of inflation affects the non-linear nature of the exchange rate dynamics.

The rest of the paper is structured as follows. In section 2 we present the Markov-switching model and discuss some of its features. In section 3 we describe the estimation process and the data used, and in section 4 we present the results. Next, in section 5 we analyse the implications of our results for exchange rate modelling and then present in section 6 a non-linear model that is capable of explaining our empirical findings.

2 The model

The non-linear model we consider is derived from the Markov-switching autoregressive (MS-AR) models popularised by Hamilton (1989) as a way of characterising expansions and contractions in empirical business cycle research. The MS-AR framework can be readily extended to various settings. In our analysis, we use the Markov-switching model to detect switches in the regressors and/or intercept of the benchmark model presented in Mark (1995) which includes a direct relationship between the (log of the) exchange rate and its (log) fundamental value. More specifically, our model is written as:

$$\Delta e_t = e_t - e_{t-1} = \alpha_{s_t} - \beta z_{t-1} + \varepsilon_t \qquad \varepsilon_t \sim N(0, \sigma^2_{r_t})$$

Where Δe_t represents the difference of the log exchange rate in month t relative to the previous and z_{t-1} stands for the deviation of the (log of the) exchange rate from its fundamental value at time $t-1$, so:

$$z_{t-1} = e_{t-1} - f_{t-1}$$

With f_{t-1} representing the date $(t-1)$ fundamental. Hence we can write f_{t-1} as:

$$f_{t-1} = \gamma_{1,q_{t-1}} \left(p_{t-1} - p^*_{t-1}\right) + \gamma_{2,u_{t-1}} \left(i_{t-1} - i^*_{t-1}\right) + \gamma_{3,v_{t-1}} \left(m_{t-1} - m^*_{t-1}\right)$$

Whereby $m_{t-1} - m^*_{t-1}$ represents the money stock differential, $p_{t-1} - p^*_{t-1}$ the price level differential and finally $i_{t-1} - i^*_{t-1}$ the interest rate differential. Further, we postulate the existence of a number of unobserved variables that take on the value one or two. More specifically, we extend the set-up of McConnell and Perez Quiros (2000) and Dewachter (2001) and allow for up to five separate and independent latent variables (namely s_t, q_t, u_t, v_t and r_t) for the dynamics of the intercept, the regressor coefficients and the variance.[3] These variables characterise the state or regime that the process is in. We assume that the stochastic process generating these unobservable regimes is an ergodic, irreducible first order Markov chain. Hence the process for these unobserved variables is presumed to depend on past realisations of e and itself only through its first lag. To overcome the potential problem of multicollinearity between money growth and inflation differential, we estimate two variants of the above equation. More specifically, we estimate the equation once with money growth and interest rate differentials and once with inflation and interest rate differentials as explanatory variables.[4]

Note that an attractive feature of the model is that a variety of behaviour is allowed. No prior information regarding the dates or the sizes of the states is required. In particular there could be asymmetries in the persistence of the states and we do not impose that the coefficients should be either significant or insignificant.[5,6]

[3] We postulate a separate and independent latent variable for the variance since we want to avoid that switches in the variances regime would influence our results for switches in the intercept or slope coefficients.

[4] Note that since the results were qualitatively the same, we only report in the paper the results for the regression with inflation and interest rate differentials as explanatory variables.

[5] As an alternative to our specification, we could have made use of a TAR/STAR/ESTAR model (see for instance Kilian and Taylor, 2003). Such model specifications are based on the view that the nominal exchange rate reverts back to its ppp-value at least in the long run (or in other words that the real exchange rate is mean reverting). In our paper, however, we prefer to take a more agnostic approach. We start out by analysing the possible non-linear nature between the change in the nominal exchange rate and its underlying fundamentals, whereby the fundamentals include beyond inflation differentials also interest rate differentials and money growth differentials. Hence, we make use of a rather general model that does not impose any a priori relationship between the exchange rate and its underlying fundamental. In addition, the model allows for non-linear dynamics between the exchange rate and each of its underlying fundamentals separately (as each of the coefficients is governed by a separate and independent latent variable). As such the approach is less restrictive than the TAR/STAR/ESTAR approach.

[6] Another alternative approach woudl be to estimate a model where the transition probabilities of the

3 Estimation process and the variables used

3.1 The variables

To estimate the aforementioned models, we choose to work with monthly data on the exchange rates and various fundamentals. For the low inflation countries, data on the home currency price for the exchange rate, the money supply, the price level and the domestic interest rate was obtained for Germany, France, Italy, Japan, the United Kingdom and the United States. The data series for the low inflation countries has been derived from the BIS except for the 10-year government bond yields which are obtained from the Global Financial Market database.

For the high inflation countries, data on the same variables was obtained for Argentina, Bolivia, Brazil, Chile, Columbia and Ecuador.[7] Here IMF data was only used if the national statistical offices did not provide the data. It should be noted that for all country data consistency with the US data was ensured.

As regards the exchange rate data, while we use for low inflation countries the standard official exchange rate series, for high inflation countries, we complement our analysis by also using market-determined (also known as 'black 'market or parallel) exchange rates as provided by Reinhart-Rogoff (2004). Using the parallel exchange rate could be beneficial since it has the advantage of being determined in a free market, and hence is not obviously contaminated by the distortionary effects of government policy.

Finally, we also need to determine the sample period over which the exchange rate in the high inflation countries studied is floating. For this, we use as a starting point the exchange rate classification as presented in Reinhart and Rogoff (2004) but extend the analysis with information from the Latin American Development Bank.

3.2 The Estimation process

In this paper, we follow the maximum likelihood approach to estimate the Markov-switching model. As the results from estimating the model were consistent for official and black market exchange rate series, only the official exchange rate series results are reported below.[8] In order to increase the probability that we reach the global maximum of the likelihood function, we randomized a number of different starting values and we used the estimates associated with the highest likelihood value.[9]

To test the significance of the Markov-switching model across regimes, it was pointed out by Hansen (1992) that classical test statistics are not asymptotically χ^2 distributed

Markov chain depend on the dynamics in the underlying fundamentals. Such a set-up was for instance pursued in Vansteenkiste (2006) within a similar context. Results from that analysis would show that in general, a larger deviation from fundamentals would tend to increase the probability of switching to the regime where changes in the fundamentals determine significantly the exchange rate changes. However, the problem with this approach is that first it requires (as is the case for the TAR/STAR/ESTAR models) an assumption regarding the fundamental model driving the exchange rate (be it PPP or a monetary model). Moreover, it also involves a sacrifice with regards the number of fundamentals that could be included in the regression. In practice, this would imply that for technical reasons all fundamentals would be imposed to switch at the same time, which is a restriction we do not impose in this paper.

[7] For some high inflation countries, data availability was restricted. More information about the samples and data availability can be found in Annex A.

[8] The results for the black market exchange rate series are available upon request from the authors.

[9] Due to the computational complexity of the model, this number has been set to 30.

in this case. These test statistics are all based on regularity conditions ensuring that the likelihood surface is locally quadratic and that the score-vector has a non-zero variance. These conditions are violated in the case of a Markov-switching model. Hence, the use of the standard distribution would therefore cause a bias of the test against the null. To circumvent this problem, Hansen (1992) has proposed an alternative likelihood ratio test, in which empirical process theory is used to bound the asymptotic distribution of a suitable standardised likelihood ratio statistic, which is applicable when the assumptions of standard theory are violated. For the various regressions performed, the grid range and the size of the step have been made dependent on the regression outcome. For the probabilities, the grid always ranged between $[0.001, 0.991]$ in steps of 0.11.

4 The results

4.1 The low inflation countries

Table 1: Hansen LR Test Results for Low Inflation Countries

	Germany	France	Italy	UK	Japan
Switching in intercept and slope					
$\alpha_0 = \alpha_1$	0.2	0.1	0.9	0.1	0.7
$\gamma_{10} = \gamma_{11}$	2.1	6.1*	7.2*	9.3*	4.5*
$\gamma_{20} = \gamma_{21}$	5.4	7.6	2.1	6.6	5.2
Switching in intercept					
$\alpha_0 = \alpha_1$	0.6	0.5	0.5	0.9	0.2
Switching in slope					
$\gamma_{10} = \gamma_{11}$	11.2*	5.4*	6.7*	7.8*	9.4*
$\gamma_{20} = \gamma_{21}$	10.3*	5.5*	6.2*	9.7*	6.6*

*A * indicates a rejection of the null at a 5% significance level. Equation:* $e_t - e_{t-1} = \alpha_i + \beta\left[e_{t-1} + \gamma_{1j}(p_{t-1} - p^*_{t-1}) + \gamma_{2k}(GBY_{t-1} - GBY^*_{t-1})\right]$ where $\varepsilon_t \sim N(0, \sigma^2_m)$, i, j, k, l, m *=0 or 1.* Further *e represents the (log of the) exchange rate, p the price level and GBY stands for the government bond yield. Note that the switches in the slope and volatility are not constrained to occur at the same moment in time.*

Table 1 shows the Hansen LR tests for the low inflation countries.[10] As will be remembered the Hansen LR test allows us to test for the equality of the intercepts and the slopes in the different regimes identified by the Markov-switching model. We have considered three scenarios for the regime switches. In the first one we test whether there are switches in the intercept and the slope, in the second case we allow for switches in the intercept, and in the third case we only allow for switches in the slope.

A first conclusion from table 1 is that the model identifies many significant switches in the slope coefficients. In particular, switches in the slope coefficients are almost always significant, except for the money stock differential.

Table 2 presents the estimates of the intercepts and slope coefficients obtained in the different regimes. The most remarkable result is that the slope coefficients often switch

[10] For Italy, France and the United Kingdom the regression were also run including a German fundamental given that they were part of ERM for some part of the sample period. The German fundamental (i.e. German call money market rate or German M3) was however for all countries but Italy not significant. For this reason, we prefer to present the regression results without the German fundamental.

Table 2: Estimates Fit to Individual Low Inflation Countries (74:8-98:11)

Parameter	Germany	France	Italy	UK	Japan
α	-0.938**	-2.547**	-13.651**	3.146**	-4.872**
	0.066	0.316	4.212	1.128	0.326
β	-0.038**	-0.121**	-0.002*	-0.002**	-0.008**
	0.014	0.024	0.001	0.000	0.002
γ_{10}	-2.687	-2.827**	3.829	-1.948	5.368
	1.666	0.770	8.853	2.149	4.882
γ_{11}	-0.491**	3.177	-2.020**	-8.867**	9.669**
	0.080	2.873	1.027	1.983	1.742
γ_{20}	0.233**	0.079	1.416**	0.000	0.588**
	0.104	0.070	0.313	0.233	0.215
γ_{21}	0.107	0.079**	-6.384	1.841**	3.071
	0.058	0.013	4.639	0.345	2.643
σ_0	0.005**	0.004**	0.001**	0.000**	0.002**
	0.001	0.001	0.001	0.000	0.000
σ_1	0.000**	0.000**	0.000**	0.003**	0.004**
	0.002	0.002	0.004	0.001	0.001
q_{00}	0.730	0.991	0.783	0.457	0.996
q_{11}	0.666	0.977	0.777	0.911	0.344
u_{00}	0.881	0.488	0.958	0.759	0.949
u_{11}	0.704	0.985	0.633	0.291	0.525
r_{00}	0.380	0.868	0.954	0.862	0.573
r_{11}	0.594	0.877	0.947	0.987	0.696

*Note: * denotes significance at a 10% level, ** at a 5% level. Equation:* $e_t - e_{t-1} = \alpha + \beta\left[e_{t-1} + \gamma_{1j}(p_{t-1} - p^*_{t-1}) + \gamma_{2k}(GBY_{t-1} - GBY^*_{t-1})\right]$ where $\varepsilon_t \sim N(0, \sigma_i^2)$, $i, j, k, l = 0$ *or 1.* Further e *represents the (log of the) exchange rate,* p *the price level,* M *the money stock nd* GBY *stands for the government bond yield. Note that the switches in the slope and volatility are not constrained to occur at the same moment in time.* q_{00}/q_{11}, u_{00}/u_{11}, r_{00}/r_{11} *are the transition probabilities for the price level, interest rate, and volatility regimes respectively.*

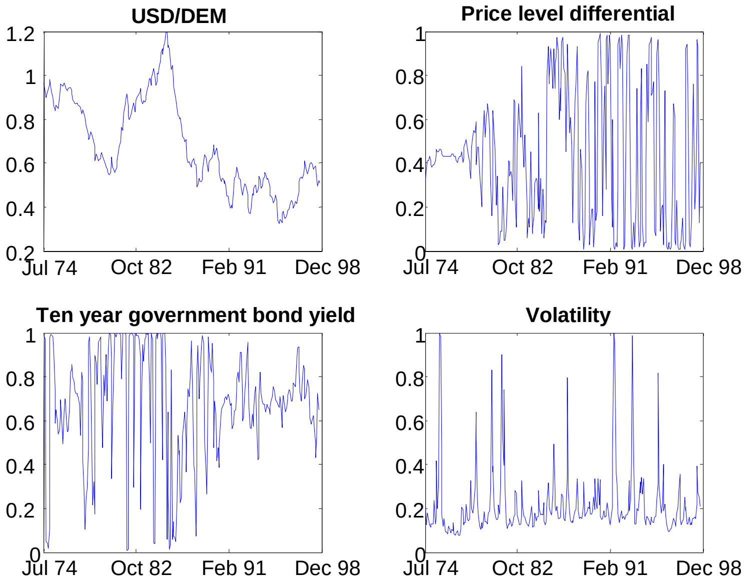

Figure 1: Smoothed Probabilities for Coefficient and Volatility Switches for German Data

between significant and non-significant values (with the exception of the coefficients of the relative money supply which are in most cases never significant), suggesting that in one regime the variables in question have a significant effect on the exchange rate, while in the other regime their effect is not significantly different from zero. Moreover, it is often the case that if the slope coefficient of one fundamental is significant in the first regime, this implies significance of the other fundamental(s) in the second regime and vice versa.

4.2 The High Inflation Countries

How do these results compare with the results obtained for the high inflation countries? Tables 3 and 4 give an answer to this question. In table 3 we present the Hansen LR tests for the significance of the switches in regimes (intercepts and slopes) in the high inflation countries. The contrast with the low inflation countries is striking. We find significant regime switches in all countries, but these switches are never due to switches in the slope. They are caused (if they happen at all) exclusively by switches in the intercept. Thus in the high inflation countries, there have been switches in the average level of appreciation, but the explanatory power of the independent variables (inflation, money supply, interest rate) has remained unchanged. This result contrasts with the results of the low inflation countries in which the explanatory power of these independent variables appears to switch frequently.

In table 4 we show the intercept and slopes in the different regimes for the high inflation countries. We observe that, in contrast to the low inflation countries, the slope coefficients are always significant and that often the switches only occur between two

Table 3: Hansen LR Test Results for High Inflation Countries

	Argentina	Bolivia	Brazil	Chile	Ecuador
Switching in intercept and slope					
$\alpha_0 = \alpha_1$	6.5*	4.7*	7.8*	10.2*	9.3*
$\gamma_{10} = \gamma_{11}$	0.3	0.4	0.9	0.6	1.6
$\gamma_{20} = \gamma_{21}$	1.1	1.3	0.1	-	0.2
Switching in intercept					
$\alpha_0 = \alpha_1$	18.4	14.5	9.5	15.4*	5.6
Switching in slope					
$\gamma_{10} = \gamma_{11}$	0.2	0.4	0.8	0.5	0.7
$\gamma_{20} = \gamma_{21}$	0.4	0.9	0.2	-	0.4

A * indicates whether the LR value found indicates a rejection of the null at a 5% significance level. Equation: $e_t - e_{t-1} = \alpha + \beta\left[e_{t-1} + \gamma_1(p_{t-1} - p^*_{t-1}) + \gamma_2(LR_{t-1} - LR^*_{t-1})\right]$ where $\varepsilon_t \sim N(0, \sigma_i^2)$, i and *j =0 or 1, where e represents the log of the official exchange rate, p the price level and LR stands for the lending rate. Note that no constraints were imposed on the timing, values or significance of the regime switches.*

Table 4: Estimates Fit to High Inflation Countries (With the Official Exchange Rate - Long Sample)

Parameter	**Argentina**	**Bolivia**	**Brazil**	**Chile**	**Ecuador**
α_0	1.057**	2.224**	-2.343**	7.583**	6.969**
	0.406	0.162	0.957	1.013	0.254
α_1	0.834**	5.355**	-4.273**	3.210**	3.368**
	0.350	0.313	0.053	0.848	1.501
β	-0.002**	-0.325**	-0.391**	-0.025**	-0.021**
	0.001	0.108	0.071	0.007	0.006
γ_1	-0.068**	-1.435**	-1.004**	-0.889**	-1.048**
	0.016	0.210	0.035	0.068	0.051
γ_2	0.786**	-1.816**	-0.666**	0.341**	-5.009**
	0.223	0.625	0.138	0.040	0.664
σ_0	0.250**	0.647**	0.014**	0.001**	0.074**
	0.002	0.079	0.020	0.572	0.012
σ_1	0.001**	0.479**	0.021**	8.575**	0.016**
	0.000	0.078	0.002	4.243	0.001
s_{00}	0.858	0.984	0.939	0.715	0.982
s_{11}	0.928	0.996	0.853	0.933	0.000
r_{00}	0.703	0.793	0.493	0.476	0.555
r_{11}	0.869	0.724	0.937	0.377	0.903

Note: standard errors are in parentheses, * denotes significance at a 10% level, ** at a 5% level. Equation: $e_t - e_{t-1} = \alpha + \beta\left[e_{t-1} + \gamma_1(p_{t-1} - p^*_{t-1}) + \gamma_2(LR_{t-1} - LR^*_{t-1})\right]$ where $\varepsilon_t \sim N(0, \sigma_i^2)$, i and *j =0 or 1, where e represents the log of the official exchange rate, p the price level, and LR stands for the lending rate. Note that no constraints were imposed on the timing, values or significance of the regime switches.* s_{00}/s_{11}, r_{00}/r_{11} *are the transition probabilities for the intercept and volatility regimes respectively.*

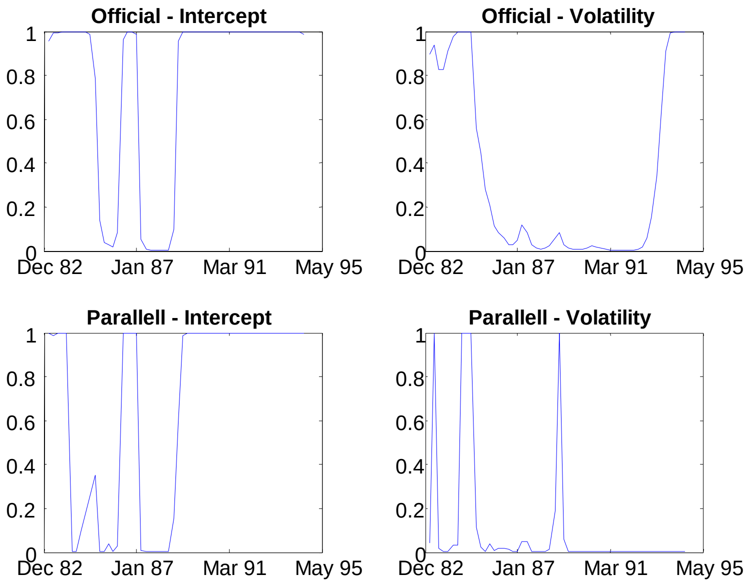

Figure 2: Smoothed Probabilities for Intercept and Volatility Switches (Brazilian Data)

significant intercepts.[11] Hence, despite the more drastic economic and political changes experienced in the high inflation countries, the Hansen tests in table 3 do not support the existence of statistically significant structural changes on the relationship between exchange rates and fundamentals of these economies. In other words none of the changes on the macreconomic and political environment seem to have forced a truly statistical change in the relationship between exchange rates and fundamentals.

[11]In table 4 we only report the Markov-switching Model that allows for switches in the intercept but not in the slope. This is our preferred specification based on the Hansen LR test statistics in table 3. However, it may be important to remark that under a Markov-switching Model with both switches in the intercept and the slope, all coefficients estimated were always significant, hence confirming our stated conclusion.

5 Theoretical Issues

The results discussed in the previous section can be summarised as follows. The relation between the exchange rate and the fundamentals of low inflation countries is characterised by frequent regime shifts. We found that the coefficients of these fundamentals change over time quite often from significant values to insignificant ones, and vice versa. This feature is absent in the exchange rate equations of high inflation countries. In those countries we find that the coefficients of the fundamentals are quite stable (only the intercept switches).

These results suggest that for the high inflation countries the linear first generation model may be the right framework for explaining the movements of these countries' exchange rates. This is not the case for the low inflation countries, whose exchange rates cannot be explained by a stable linear relation with underlying fundamentals.

Any explanation of these empirical results must be capable of accounting for the differences observed in the stability of the exchange rate equations between low and high inflation countries. There are two alternative explanations. The first alternative is based on the second-generation model. We claim that this explanation is unsatisfactory. The second-generation model is based on explicit utility maximisation of a representative agent. In this model the structural instability of the coefficients in the exchange rate equations can be explained by shifts in the underlying stochastic structure, which may or may not be induced by changes in policy regimes. The contrasting evidence between high and low inflation countries, however, makes this explanation implausible. If anything, high inflation countries experience stronger changes in the underlying stochastic structure (mainly induced by shifts in policy regimes) than low inflation countries. And yet it is in the high inflation countries that the linear first generation model seems to be doing well while it fails for the low inflation countries.

For this reason our preferred explanation is based on nonlinearities. In what follows, we outline the nature of non-linear features that in our view are capable of explaining the unstable relation between the exchange rate and its underlying fundamentals in low inflation countries. In this section we only briefly sketch the nature of these nonlinearities and how these affect exchange rate models. In the next section we present a simple model formalising some of these ideas.

An important nonlinearity has been stressed by Obstfeld and Rogoff (2000) who show that many of the current puzzles in international macroeconomics can be explained by transactions costs in the goods markets (see also Dumas (1992), Uppal, Sercu and Van Hulle (1995),and De Grauwe and Grimaldi (2005)). Transactions costs in the goods markets create a band of inaction within which international price differentials are not arbitraged away; only price differentials exceeding transactions costs (outside the band) are profitable to arbitrage. Transactions costs must be overcome for trade in goods to take place. To illustrate the magnitude of these transactions costs, De Grauwe and Grimaldi (2005) show the price dispersion of a sample of exactly the same products within the European Union. We observe price differentials of up to 40%. This suggests that producers price to market. Such pricing strategies can however only be applied successfully if transactions costs prevent arbitrage. Thus, these large price differentials suggest that transactions costs for traded goods are large, with the order of magnitude being between 20 and 40%.[12] In our case, introducing these transactions costs can contribute to understanding the difference in the relationship between the exchange rate and its fundamentals for low and high inflation countries.[13] To see this, consider the following set-up.

The existence of transactions costs (say as a fixed proportion of the prices of products) defines a band in which arbitrage relations, such as the PPP relation, do not hold. This is the case in both the low and high inflation countries. Now introduce exogenous shocks in the underlying fundamental values of the exchange rate. In the low inflation countries, many shocks tend to be small relative to the transactions cost band (e.g. differential inflation shocks are typically 1 or 2% per year). Hence, arbitrage will not be profitable in these cases and will remain absent. Some shocks, however, are large relative to the transactions cost band implying that arbitrage will take place. As a consequence, the relation between exchange rates and their underlying fundamentals will be unstable. In contrast, in the high inflation countries, shocks in the fundamentals (especially nominal shocks) tend to be large relative to the transactions costs band, imposing strong arbitrage relations. This implies that the relation between the exchange rate and its fundamentals remains stable. The empirical importance of these transactions costs have also been confirmed empirically, for instance in Peel, Sarno and Taylor (2001) and Kilian and Taylor (2003).

As stressed earlier, this is only a broad sketch of nonlinearities in exchange rate models capable of explaining the results obtained in this paper. In the next section we present a simple non-linear model that allows us to capture some of the general ideas developed in this section.

6 A simple non-linear model with transactions costs

In this section, we develop a non-linear model that is as parsimonious as possible.

The exchange rate is e_t and its fundamental value is represented by f_t. The latter could be the price level, or more generally a vector of variables that determine the equilibrium value of the exchange rate. We assume that it is driven by a random walk process, i.e.:

$$f_t = f_{t-1} + \varepsilon_t \tag{1}$$

We assume fixed transactions costs, τ. The effect of these transactions costs is to prevent goods arbitrage. As a result, as long as the exchange rate is within its transactions cost band, there is no mechanism that drives the exchange rate towards its fundamental value. More formally we postulate the following process: If,

$$|e_t - f_t| < \tau, \qquad e_t - e_{t-1} = \eta_t \tag{2}$$

where η_t is a white noise variable. If,

$$|e_t - f_t| > \tau, \qquad e_t - e_{t-1} = \vartheta(f_{t-1} - e_{t-1}) + \eta_t \tag{3}$$

[12] It can be argued that for non-traded goods transaction costs are even higher (see Obstfeld and Rogoff, 2000).

[13] Transaction costs also exist in financial markets. However, here there is less consensus of the role of these transaction costs. In addition the magnitudes are significantly smaller. Nevertheless, several studies indicate the importance of such transaction costs to account for the failure of for instance in the uncovered interest parity (see for instance Chang, Reed, Lei and Rhee, 2003).

In words, when the difference between the exchange rate and its fundamental value is within the transactions cost band given by τ, the changes in the exchange rate are white noise. When the difference between the exchange rate and its fundamental value is larger than the fixed transactions costs, the exchange rate tends to return to its fundamental value. The speed with which this happens is determined by the parameter ϑ. In rational expectations models this parameter will typically be influenced by the structural parameters of the model, including the speed of adjustment in the goods market.[14]

Equations (1) to (3) present a very simple non-linear model of the exchange rate. In order to judge its empirical relevance we simulate the above presented model and use it to analyse whether the model is capable of replicating some of the empirical features analysed in the previous sections.[15] We will assume different values of the speed of adjustment parameter ϑ and of the transactions cost parameter τ. We then apply the Markov-switching and time-varying parameter methodology to analyse under what conditions this simple model produces regime switches that are similar to those detected in the data.

We considered cases that come close to representing the situations of low and high inflation countries. More specifically, low inflation countries are those for which the transactions cost band is high compared to the size of the shocks in the fundamentals. In addition we assume that in these countries the speed of adjustment of prices is low. This is the case represented by $\vartheta = -0.02/\ \tau = 2$. In high inflation countries the size of the transactions cost band is low compared to the size of the shocks in the fundamentals, and the speed of adjustment of prices is high. This is the case represented by $\vartheta = -0.04/\ \tau = 1$[16]. The parameters have been mainly based on findings from the empirical literature.

As regards the speed of adjustment parameter, there exists a vast literature on the determination and estimation of the half life of deviations from PPP. In summarizing the results from studies using long-horizon data, Froot and Rogoff (1995) and Rogoff (1996) report the current consensus in the literature that the half-life of a shock (the time it takes for the shock to dissipate by 50 percent) to the real exchange rate is about three to five years, implying a slow parity reversion rate of between 13 to 20 percent per year.[17] However, in this paper we are more interested in the speed of convergence of the

[14] The fact that the parameter ϑ is not infinite implies thus that there are some price rigidities that do not come from transactions costs.

[15] To simulate the model we assume that the fundamental consists of its past observations plus a random shock which is assumed to be Gaussian. We generate 300 such random shocks for each iteration. In total we run the simulation 100 times.

[16] We follow this procedure of setting the parameter τ (transaction cost band) different for low and high inflation countries because in the simulations we assume the same variance of the shocks in the fundamentals. Alternatively we could have assumed different variances (high for high inflation and low for low inflation countries) and the same values τ. The two procedures yield qualitatively the same results. What matters is that in the high inflation countries the width of the transaction cost band is low relative to the size of the shocks. The opposite holds for the low inflation countries.

[17] Abuaf and Jorion (1990) use data on bilateral real exchange rates between the United States and several industrial countries during the twentieth century, and find average half-lives of deviations from parity of a little over three years. Frankel (1986) and Lothian and Taylor (1996) use two centuries of annual data on the sterling-dollar real exchange rate in calculating half-lives of about five years. Wu (1996) and Papell (1997) use panel data methods on quarterly post–Bretton Woods data to derive half-lives of between two to three years.

exchange rate towards a more broadly defined fundamental value. Here little estimates are available. Looking at the estimation results from Mark (1995) we find that half life convergence would take around 4 years for low inflation countries. Based on this information, we select a speed of adjustment with a half life of 2 years, which is on the lower bound of existing studies and a conservative input for our purpose (implying that ϑ equals -0.02 when assuming we use monthly observations). For high inflation countries, few studies are available (even to determine the half life of deviations from PPP), however the consensus is that the PPP relation holds much tighter in high than in low inflation countries (see Frenkel, 1978 and Chinn, 2001). González Anaya (2000) for instance finds half lives during the high inflation episodes in Latin American countries of 6 months to 2 years. This result is further confirmed by Yazgan (2003) who finds a half life of one year for Turkey. Hence, we set ϑ equal to -0.04 implying a half life of one year.

As regards the transactions cost parameter, it was already stressed in the previous section that these tend to be particularly large in the goods market of low inflation, where observed price differentials suggest transactions costs for traded goods as large as 20 to even 40% whereas the size of the shocks to the underlying fundamentals are typically only a few percentage points per annum. Hence for low inflation countries, this would suggest that the exchange rate would hardly ever leave the transactions cost band. For our simulation exercise we opt for a parameter of $\tau = 2$ which would imply that for 96% of shocks to fundamentals the exchange rate would not leave the transactions cost band. By contrast for high inflation countries, we set $\tau = 1$ which means that for 32% of the shocks to fundamentals exceed the transactions cost band.

The results are shown in tables 5 and 6. Our results are quite interesting. We find that the simple non-linear model predicts that in low inflation countries there are frequent switches in regimes, i.e. the slope coefficients of the fundamental variables switches regularly. No such regime switches in the slope coefficients are observed for the high inflation countries. Similar results were obtained when estimating the time-varying parameter model. Here on average we found that for 42 observations out of 100 the slope coefficient was significantly different from zero for low inflation countries whereas for high inflation countries it was on average significantly different from zero in 98 observations out of 100.

Table 5: Hansen LR Test Results, Number of Simulations with Null Rejected at 5 Percent Level

Switches in	intercept, slope and volatility			intercept	slope
$H_o:$	$\alpha_0 = \alpha_1$	$\gamma_0 = \gamma_1$	$\sigma_0 = \sigma_1$	$\alpha_0 = \alpha_1$	$\gamma_0 = \gamma_1$
$\vartheta = -0.02/ \ \tau = 2$ "Low Inflation Country"					
	2	92	76	5	96
$\vartheta = -0.04/ \ \tau = 1$ "High Inflation Country"					
	87	22	65	92	3

100 simulations with 300 observations each were run. For each simulation the Markov-switching model was next estimated and then the Hansen test was estimated. The estimated Markov-switching model is $e_t - e_{t-1} = \alpha_i + \beta\left[\gamma_{1j}(fund_{t-1} - e_{t-1})\right] + \varepsilon_t$ where $\varepsilon_t \sim N(0, \sigma_k^2)$, i, j, k *=0 or 1.*

Table 6: Estimates fit to Non-linear Model, Number of Simulations with Significant Coefficient at 10 Percent Level

	α_0	α_1	γ_0	γ_1
$\vartheta = -0.02/\ \tau = 2$ "Low Inflation Country"	23	–	76	21
$\vartheta = -0.04/\ \tau = 1$ "High Inflation Country"	98	76	100	–

Note: The results in the table shows the number of simulations with significant coefficients (at 10% level) out of 100 simulations with 300 observations each were run. The estimated Markov-switching model for each simulation for the low inflation countries is $e_t - e_{t-1} = \alpha + \beta\left[\gamma_{1j}(fund_{t-1} - e_{t-1})\right] + \varepsilon_t$ where $\varepsilon_t \sim N(0, \sigma_k^2)$, j, k *=0 or 1 and for high inflation countries it is* $e_t - e_{t-1} = \alpha_j + \beta\left[\gamma_1(fund_{t-1} - e_{t-1})\right] + \varepsilon_t$ where $\varepsilon_t \sim N(0, \sigma_k^2)$, j, k =0 or 1. *If in one of the two regimes the coefficient was significant, it was systematically classified as regime 0.*

7 Concluding Remarks

Characterising the nature of the relationship between exchange rate changes and the changes in its underlying fundamentals has long been an objective of empirical international macroeconomics. Although this research has contributed to our understanding of the behaviour of the exchange rates, it is also true that this empirical research has been unable to validate the existing theoretical models. In particular, the 'first generation models' of the exchange rates that were developed during the 1970s have been rejected at least when using data of the major industrial countries. The 'second generation models' based on explicit utility maximisation of agents have not produced sharp enough testable propositions allowing for their refutation by the data. As a result, they have neither been confirmed nor refuted.

In this paper, we tested whether the relationship between the changes in the nominal exchange rate and the news in its underlying fundamentals has non-linear features. In order to do so, we developed a Markov-switching and applied it to a sample of low inflation and high inflation countries.

The empirical analysis shows that for the high inflation countries the first generation models appear to work well: the relationship between news in the fundamentals and exchange rate changes is stable and always significant. This is not the case, however, for the low inflation countries, where frequent regime switches occur.

We developed a non-linear model that is capable of explaining our empirical findings. The model is based on the existence of transactions costs in the goods markets. We found that this simple non-linear model is capable of replicating the empirical evidence uncovered in this paper. More specifically the model predicts that in countries where shocks in fundamentals are low in comparison with the transactions cost band (low inflation countries), frequent regime switches in the link between the exchange rate and its fundamentals must occur. This is not the case in high inflation countries where the size of the shocks in fundamentals is large relative to the transactions cost band.

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A Data Availability and Sample Selected

The countries included in the analysis are: Argentina, Bolivia, Brazil, Chile Colombia, Ecuador, France, Germany, Italy, Japan and the UK. Information on the home currency-dollar exchange rate and five fundamentals was retrieved on a monthly and quarterly basis. More specifically, this set of fundamentals covers:

1. The price level for the country concerned, defined as the consumer price level.

2. The money supply for the country under scrutiny, for all low inflation countries this is M3 while for high inflation countries this is M0.

3. The money market rate, which is used as a measure of the short-term interest rate for high inflation countries

4. The interbank rate, which is used as a measure of he short-term interest rate for low inflation countries.

5. The lending rate and the long-term government bond yield which are both proxies of the long-term interest rate. The latter was however only available for the low inflation countries

In table 7 below, the sample periods for the regressions are reported for the monthly data. For the quarterly observations, the same time periods were available then the figures were transformed to quarters rather than months.

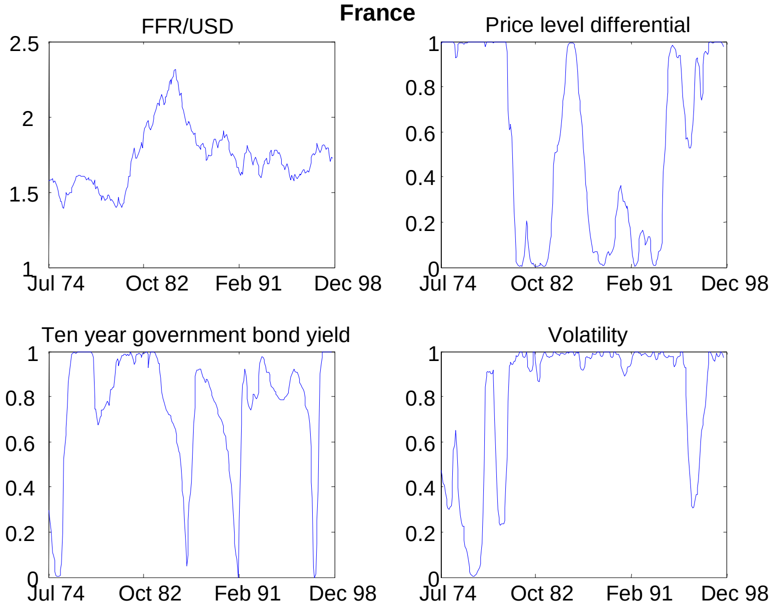
Table 7: Sample Periods Used for Estimations

	Short Sample	**Long Sample**
Low Inflation Countries		
Germany	1973M1-1998M12	
France	1973M1-1998M12	
Italy	1973M1-1998M12	
Japan	1988M1-2005M7	
United Kingdom	1973M1-2005M7	
High Inflation Countries		
Argentina	1981M3-1985M6	1979M3-1991M3
Bolivia	1980M12-1985M12	1980M12-1986M12
Brazil	1989M4-1994M5	1982M12-1994M7
Chile	1971M7-1976M6	1962M4-1978M12
Ecuador	1983M1-1998M11	

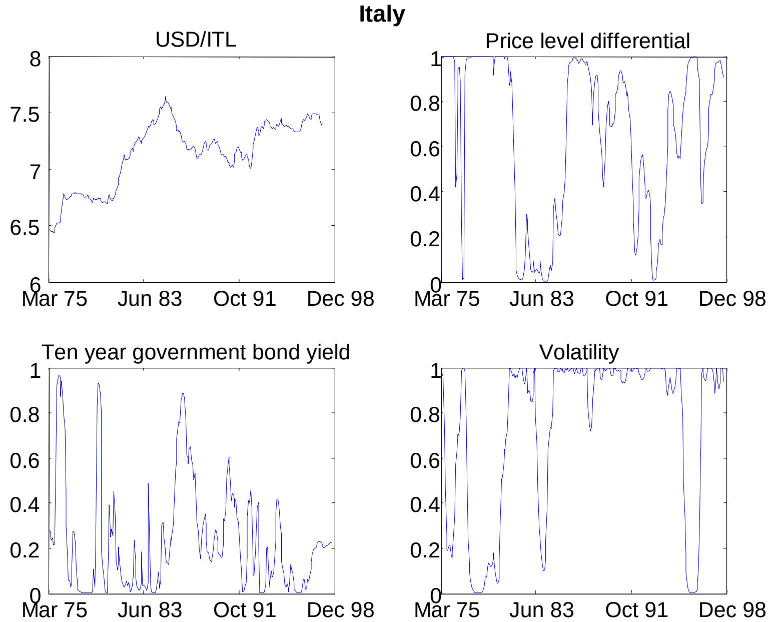
The short samples for the high inflation countries are those where the exchange rate can be defined as floating. The long sample is the full sample for which we have data for the exchange rate of the country and all its fundamentals (except Chile where only information is available for inflation). During the full sample period various exchange rate regimes could happen though. In Ecuador the exchange rate was for the full sample for which we have data never floating.

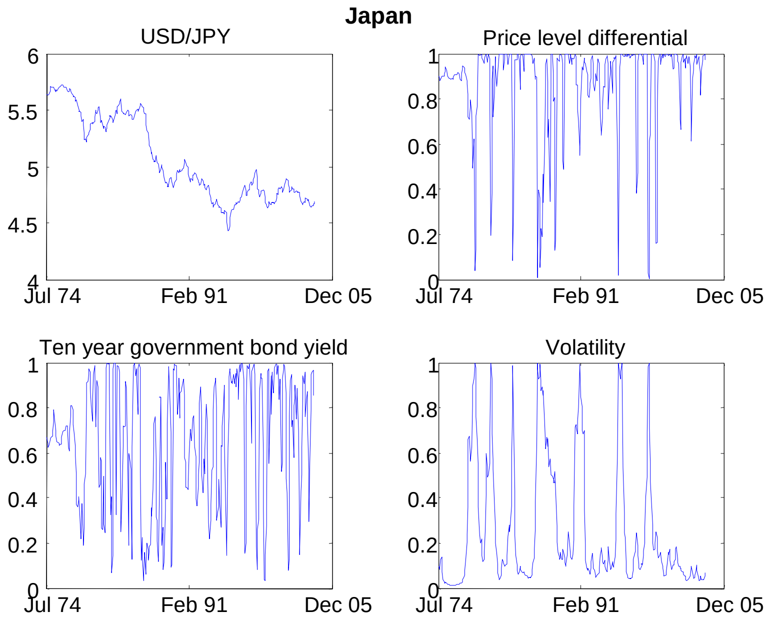
B The Smoothed Probabilities for Low Inflation Countries

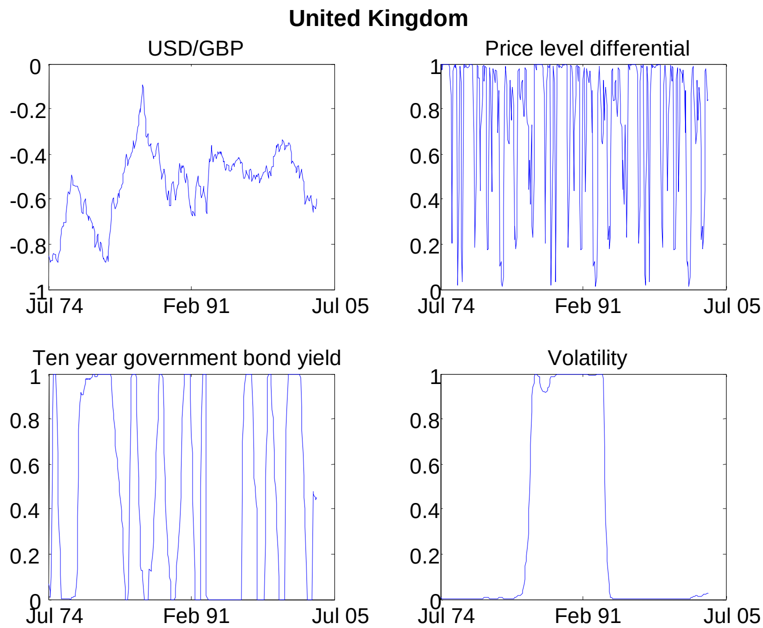
France



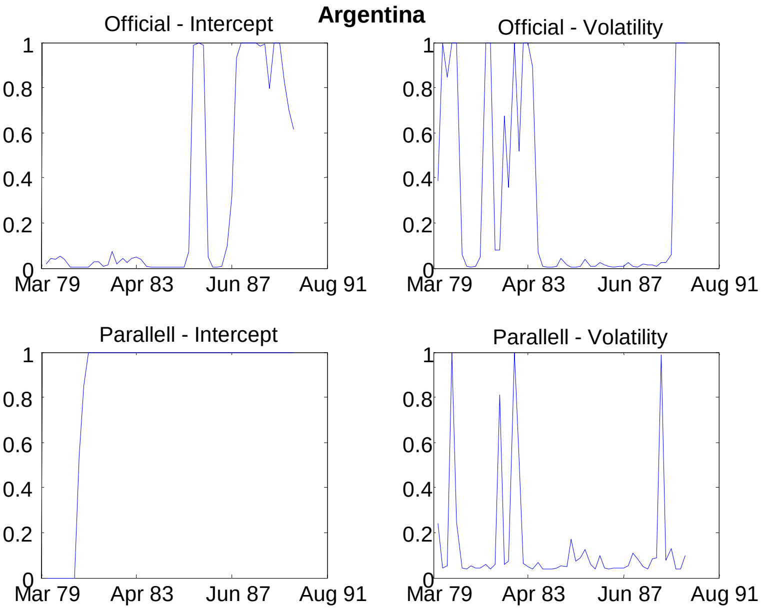
Italy

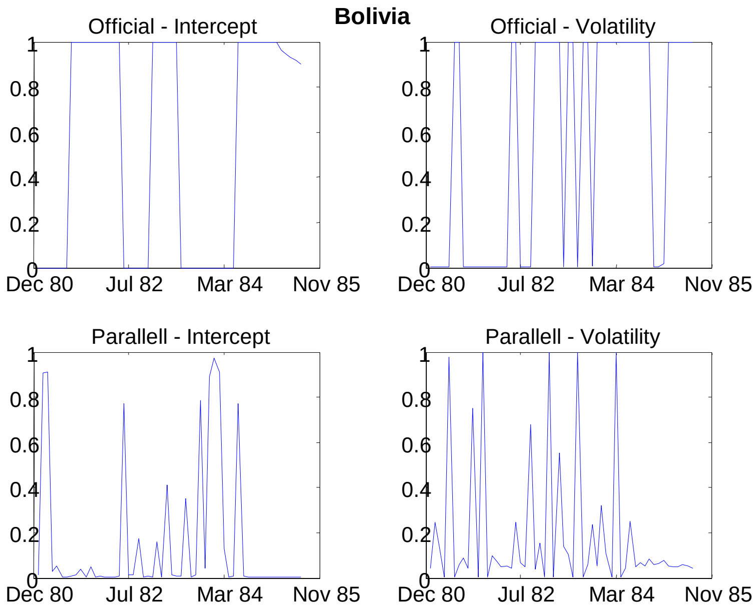


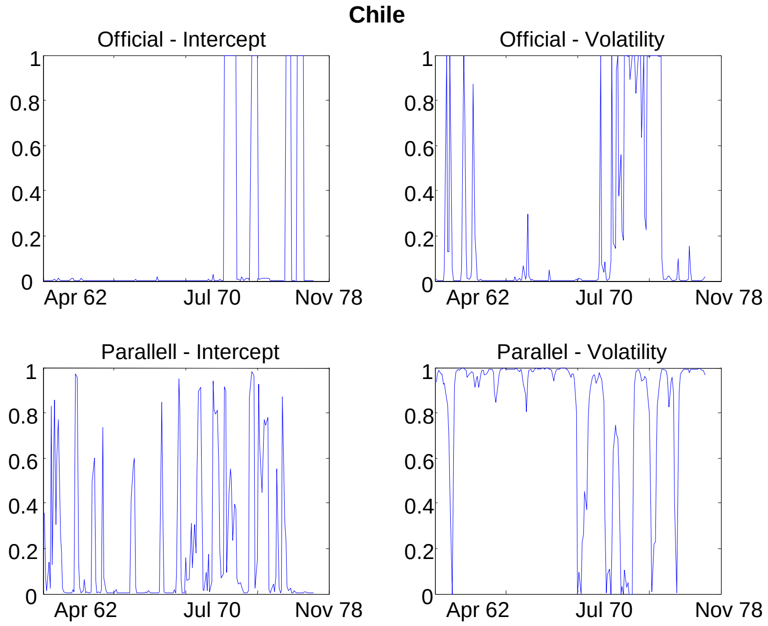

Japan
USD/JPY
6
5.5
5
4.5
4
Price level differential
1
0.8
0.6
0.4
0.2
0
Ten year government bond yield
Volatility
Jul 74
Feb 91
Dec 05

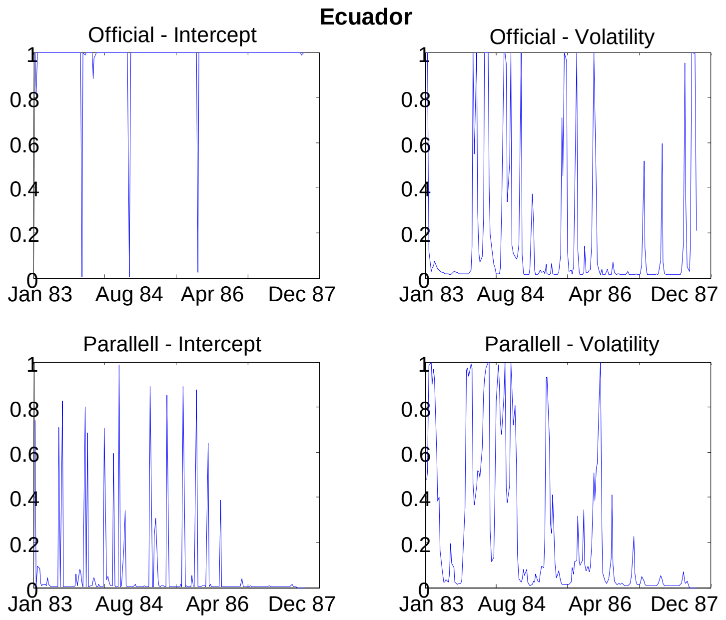

United Kingdom
USD/GBP
0
-0.2
-0.4
-0.6
-0.8
-1
Price level differential
1
0.8
0.6
0.4
0.2
0
Ten year government bond yield
Volatility
Jul 74
Feb 91
Jul 05

C The Smoothed Probabilities for high inflation countries






Chile
Official - Intercept
Official - Volatility
Parallell - Intercept
Parallel - Volatility
1
0.8
0.6
0.4
0.2
0
Apr 62
Jul 70
Nov 78


Ecuador
Official - Intercept
Official - Volatility
Parallell - Intercept
Parallell - Volatility
1
0.8
0.6
0.4
0.2
0
Jan 83
Aug 84
Apr 86
Dec 87