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Leaning against the parity

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Leaning Against the Parity

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Abstract

The paper presents evidence that the simultaneous relationship between uncovered interest rate parity (UIP) and a monetary policy function can explain the empirical failure of the former. Using the model proposed by McCallum (1994), we carry out tests for a sample of developed and emerging markets from 1995M5 to 2004M3. The results lend strong support to the view that monetary policy affects the equilibrium nominal interest rate differential between emerging economies and the US. Slow adjustment in interest rates and reaction against price changes seem to be the prominent features of the reaction function. Shocks have an asymmetric impact on the volatility of the differentials which is also significant to explain monetary policy. Finally, the dynamic properties of uncovered interest rate parity *ex post* deviations, also interpreted as risk premium, influence the equilibrium nominal interest rate differentials.

JEL classification: E43, E52, E58, F31.

Keywords: Uncovered Interest Rate Parity, Monetary Policy, Reaction Function, Interest Rates.

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1. Introduction

The uncovered interest rate parity hypothesis (UIP), under the assumption of frictionless markets, predicts that nominal interest rate differentials (*nids*) between two economies will be equal to expected changes in the exchange rate. However, the vast majority of the empirical literature has found a robust negative relationship between the two variables. Estimates suggest that an expected appreciation of the domestic currency leads to an increase in domestic interest rates, or otherwise, that an expected depreciation decreases the nominal interest rate differential (*nid*).

Flood and Rose (2001, p.3) wrote that it is always easy to motivate another look at the UIP. The reason is that UIP is used as the cornerstone of many macroeconomic models but evidence is, at best, mixed. The contrast between its widespread theoretical use and the precarious empirical support is puzzling and provides the motivation for the present paper.

Several approaches aim at explaining the UIP's empirical failure. Most of the supportive results are conditioned on particular circumstances and specific data sets[1]. Among the competing views, risk premium with time-varying components seems to be dominant. However, this explanation also presents problems which arise from the fact that risk is not directly observed [Engel (1996)].

Authors have been raising issues regarding the behaviour of monetary authorities, especially in the floating era, which might have had an impact on *nids* and, thus, on UIP deviations. For example, a related work on the subject is Calvo and Reinhart (2002). They argued that some economies suffer from the "Fear of Floating". According to them, monetary authorities of emerging countries put a heavy weight on exchange rate stabilisation when setting interest rates because their Central Banks are scared of exchange rate volatility. The "Pick your Poison" tale, put forward by Iwata and Tanner (2003), can be considered as a

[1] For example, Flood and Rose (1996), Jorion (1996), Huisman *et al.* (1998), Meredith and Chinn (1998), Bernhardsen, T. (2000), Berk and Knot (2001), Flood and Rose (2001), Muinhos *et al.* (2001), Sachsida *et al.* (2001), Francis *et al.* (2002) and Lothian and Wu (2003) used diverse approaches to explain UIRP failure.

side effect of the remedies that are used to heal an economy from the fear of floating. Their account is that monetary authorities of emerging economies have to choose between the venoms of wild exchange rate changes or increased interest rate volatility.

In a previous work, McCallum (1994a) developed a model that considers the effects of policy makers on the determination of *nids*. The paper's main idea is that the UIP's failure can be explained by the simultaneous relationship between two factors: a policy reaction function and UIP itself. The distinguishing features of the policy function is that monetary authorities slowly change interest rates, in other words, they practice "interest rate smoothing"[2], and also resist exchange rate changes, or put in another way, they "lean against the wind". The simultaneous interaction between the two factors implies an equilibrium *nid* that, under certain parameter values, can explain the empirical failure of the UIP.

Christensen (2000) tested McCallum's (1994a) idea for developed countries (the US, Germany and Japan) and, in short, he did not find supporting evidence. Our contribution to the literature is twofold. First, we apply McCallum's (1994) model for a sample of not only developed, but also emerging markets. As stressed before, there is evidence showing that the fear of floating (or resistance against exchange rate changes) is more apparent in emerging economies. In fact, we not only verify whether McCallum's (1994) approach is consistent with our data but also compare our results with Christensen's (2000) findings. Second, we check whether and how the findings change if the policy reaction function is modified to include other variables besides exchange rate changes. The monetary authorities in both McCallum's (1994a) model and Christensen's (2000) tests practice interest rate smoothing and react to exchange rate changes only, however, the general understanding is that monetary authorities also pay attention to variables such as inflation, output gap or output changes. Furthermore, it is also known that many emerging economies have recently

[2] Sack and Wieland (2000, p.205) define interest rate smoothing as being the tendency of Central Banks to alter interest rates "in sequences of small steps in the same direction and reverse the direction of interest rate movements only infrequently".

embarked into inflation target regimes. Finally, we perform estimations considering possible structural breaks. The reason is that changes in exchange rate systems, as experienced by the countries of our sample, may have implied changes in the monetary policy.

The rest of the paper is organised as follows: the next section explains the model and introduces alternative specifications for the policy reaction function; then it briefly surveys the literature that focuses on the role of monetary policy to explain *nids*; section 3 analyses the empirical results and the final section concludes.

2. The Model

UIP under the assumption of rational expectations is represented by the equation below

$$s_{t+1} - s_t = \alpha + \beta nid_t + \varepsilon_t \qquad (1)$$

where the subscript t represents time, s is the exchange rate, defined as the domestic price of the foreign currency, $nid(s)$ represents the nominal interest differential(s) between the domestic and the foreign economy, i.e. $nid_t = i_t - i_t^*$ in which i_t is the interest rate paid on a one-period bond that matures at time $t+1$ and the asterisk stands for the foreign economy; ε is the random error term, associated with the forecast of s_{t+1}, and has zero mean and unit variance. All variables, except interest rates (and, therefore, *nids*) are in logarithms. If the hypothesis regarding UIP under rational expectations holds, then, estimated parameters α and β would not be statistically different from 0 and 1, respectively, and the error term would be white-noise. However, as discussed before, the almost unanimous finding of the empirical literature is a robust negative estimate of β [3,4]. This phenomenon is at odds with the

[3] Baillie and Bollerslev (2000) and Bekaert and Hodrik (2001), for example, are few exceptions of supportive results for UIRP. See Engel (1996) and Wang and Jones (2002) for recent surveys on the subject.

theory that explains the functioning of international financial markets on the basis of arbitrage and the assumption of no frictions.

McCallum's (1994) model[5] attempts to explain this empirical failure without abandoning either the UIP hypothesis or the assumption of rational expectations. As it will be shown, however, the model considers the existence of frictions in financial markets. The policy reaction function is given below

$$nid_t = \lambda^s (s_t - s_{t-1}) + \sigma nid_{t-1} + \zeta_t \qquad (2)$$

where λ^s stands for the degree to which a Central Bank leans against the wind, σ is the parameter representing the interest rate smoothing and the white-noise term ζ represents "random policy influences". The UIP condition is rewritten as follows

$$s_t = s_{t+1}^e - nid_t + \xi_t \qquad (3)$$

where the superscript *e* represents expectation, ξ_t stands for *ex ante* deviations from UIP, which can be due to frictions in financial markets (such as risk premium) and is autoregressive, i.e. $\xi_t = \rho\xi_{t-1} + \upsilon_t$, where υ_t is white-noise[6]. Rearrange equation (3) with the nid_t in the left-hand side and then substitute it into (2) to obtain

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[4] A great part of the estimations of β are done through tests of forward exchange rate unbiasedness [see Engel (1996)]. However, note that covered interest parity, $nid_t = f_t - s_t$, in which f_t is the price of the foreign currency at time *t* for delivery at $t+1$, imply that a test of equation (1) also corresponds to a test of foreign market efficiency. The reason is that covered interest parity substituted in (1) gives $s_{t+1} - s_t = \alpha + \beta(f_t - s_t) + \varepsilon_t$.

[5] Engel (1996) explained that his model is a particular case of the more general model of Boyer and Adams (1988).

[6] The presence of the foreign interest rate in equation (2) accounts for the fact that monetary authorities abroad may also resist exchange rate changes. Alternatively, it may indicate that domestic monetary policy cares about the level of the foreign interest rate. In any case, we will be referring to equation (2) as the policy reaction function.

$$\Delta s_{t+1}^{e} = \lambda^{s} \Delta s_{t} + \sigma nid_{t-1} + \zeta_{t} - \xi_{t} \qquad (4)$$

where Δ represents the first difference operator and $\Delta s_{t+1}^{e} = s_{t+1}^{e} - s_{t}$. Postulating a bubble-free linear solution of (4) using the method of undetermined coefficients and also considering rational expectations, the model can be solved for exchange rates:

$$\Delta s_{t} = \left(\frac{\rho - \sigma}{\lambda^{s}} \right) nid_{t-1} - \frac{1}{\lambda^{s}} \zeta_{t} + \left(\frac{1}{\lambda^{s} + \sigma - \rho} \right) \upsilon_{t} \qquad (5)$$

with the assumption that $|\rho| < 1$ and the presumption that $\rho > 0$. McCallum (1994a) thought that if $\sigma \cong 0.8$ and $\lambda^{s} \cong 0.2$ in particular, then $\left(\frac{\rho - \sigma}{\lambda^{s}} \right)$ would have a similar value and the same sign of the anomalous regression's coefficient, which is normally estimated at around -3[7].

The final result of a bond bid is a possible way to interpret equation (5). In this bid, agents use rational expectations to form forecasts about future exchange rate changes while authorities implement monetary policy reacting to variables that are considered to be relevant to their objectives. The simultaneous nature of this process leads to a result that does not corroborate the prediction of UIP. The coefficients in equation (5) show that the impact of a unit change of the predetermined variable on the exchange rate change may not be equal to one in equilibrium.

Other Policy Reaction Functions

In general, Central Banks either follow a Taylor-type rule or minimise a loss function with positive weights for both inflation and output [e.g., see Blanchard and Fischer (1989) and Sack and Wieland (2000)]. Meredith and Chinn (1998, p.10) considered that the reaction

[7] Therefore, the implicit presumption is that $\rho \cong 2$.

function of McCallum (1994a) "does not incorporate variables that are usually believed to be of concern to policymakers, such as inflation and output." McCallum (1994b) also argued that monetary authorities pay more attention to inflation and output than to exchange rate movements and, hence, a monetary function as in equation (2) represents a simplification. He has also suggested to estimate different policy reaction functions and to use diverse countries. For these reasons, we conduct several empirical tests in order to assess whether lagged *nids*, exchange rates, prices, output changes, output gap or any combination of these variables affected the *nids* of a heterogenous sample of countries. We also check whether the parameters of the estimated policy reaction function are consistent with the "parable" proposed in McCallum (1994a). In spite of the fact that one should not expect those coefficients to have the same sign and the exact magnitude that McCallum (1994a) had hypothesised, as it is emphasised in the conclusion of McCallum (1994b), we will still compare our results with his guesses. For example, the following monetary policy reaction function is the one that generates the closest results with regard to McCallum's (1994) predictions when OLS estimations are used:

$$nid_t = \lambda^p (p_t - p_{t-1}) + \sigma nid_{t-1} + \zeta_t \qquad (6)$$

where p represents price and λ^p is the degree to which monetary authorities react to price changes. By virtue of relative purchasing power parity, one can substitute inflation from equation (6) for exchange rate changes and substitute it into UIP in order to obtain the same coefficients of equation (5). This modification keeps the attractive simplicity of the original

theoretical formulation[8,9]. However, as it will be shown, a more general policy reaction function gives closer results when conditional volatility is taken into account.

The methodology adopted in this paper is based on Christensen (2000). It consists in estimating the parameters $\lambda^s, \lambda^p, \sigma$ and ρ from the structural models in order to see whether the intuition behind McCallum (1994a) corresponds to the stylised facts of the actual data. As emphasised before, we extend the work in two main directions. Firstly, we apply the model to a sample of emerging markets as well as developed economies. The second contribution is that we allow the policy reaction function to depend on other variables besides exchange rate changes and also verify whether modifications in the way that policy makers react to these variables has changed, as the countries of our sample experienced periods of fixed, managed and floating exchange rates.

Finally, it is possible to define what we understand as "leaning against the parity". The economic literature labels leaning against the wind as the attempt of monetary authorities to resist exchange rate changes. As McCallum (1994a) shows, this resistance alters the predictions of UIP. But if monetary policy is also likely to take inflation into account, then relative purchasing power parity implies that UIP results can be similarly affected. Any attempt to avoid either exchange rate or price changes will alter the results of UIP and, consequently, the real interest rate parity hypothesis. Thus, leaning against exchange rates or prices imply leaning against the parity.

Simultaneity between Arbitrageurs and Policy Makers

Kugler (2000) analysed the expectation hypothesis of the term structure of interest rates using a policy reaction function. One of his conclusions is that interest rate smoothing and

[8] Taylor (1999), for instance, lists an extensive number of policy reaction functions. The aim of this paper is not to assess all possibilities but to test a selected number of those that seem reasonable from the point of view of the theory underlying the model.

[9] The solution for a version of McCallum's (1994) model in which the policy reaction function exclusively depends on price changes is available from the author upon request.

leaning against the wind lead to a negative relationship of the spot exchange rate change and the lagged forward premium and, thus, a rejection of the UIP. Bonser-Neal *et al.* (2000) presented an UIP model in which the Federal Reserve can choose to offset shocks to exchange rates via interest rates. Meredith and Chinn (1998) developed a macroeconomic model based on McCallum (1994a) and ran regressions on data generated by stochastic simulations. Their simulations replicated the sign and magnitude of the parameters usually found in the literature using bonds of short-run maturity. They concluded that UIP fails in the short run because of risk premium shocks and endogenous monetary policy. Anker (1999) argued that the failure of UIP can be a consequence of systematic monetary-policy reaction to exchange rate changes.

A number of authors have stressed the influence of monetary policy on *nids* through several different channels. Huisman *et al*. (1998) showed that monetary uncertainty is important to explain shifts in the real interest rate. Kaminsky and Leiderman (1998) explained interest rate differentials on the grounds of a government's lack of credibility. Flood and Rose (2001) argued that deviations from UIP are a necessary condition for defending the exchange rate. Cushman and Zha (1997, p.434) stated that Central Banks of small open economies "are likely to respond quickly to foreign variables, invalidating the assumption that the interest innovations are independent." Faust and Rogers (2003) found that monetary policy shocks led to UIP deviations. Using overnight Eurocurrency deposit rates, Baillie and Osterberg (1998) reached the conclusion that Central Bank interventions impact on the risk premium and thus affect UIP. Cecchetti *et al.* (2002, p.2) quote Ball (1999) who found "that adding the exchange rate to the Taylor rule improves macroeconomic performance in a model where the exchange rate has a significant role in the transmission mechanism of structural shocks and monetary policy". To summarise, the empirical and theoretical literature reports that policy makers tend to place a significant weight on inflation

or exchange rates when setting interest rates, in other words, the influence of monetary policy on the dynamics of *nids* has been largely documented.

3. Results

3.1. Data

Data were obtained from the IMF's International Financial Statistics (IFS). The emerging markets of the sample comprise the small open-economies of Argentina, Brazil, Chile, Mexico and Turkey; the developed countries are France, Italy, Spain, the UK and Germany. We used the US as the reference economy for the calculation of the *nid*[10]. The period of the tests corresponds to the interval that spans from 1995M5 to 2004M3, hence we have 107 observations for each country. However, *nids* of France were calculated until 2002M11 and Italy until 2004M2 because of data availability on interest rates. The sample period starts in the mid 1990s for two main reasons. The first is the impact of hyperinflation on the bond weights used to build the interest rates series of emerging economies which, according to the way that Central Banks calculate short term interest rates, led to serious problems of measurement error. Because of this, and the intractability of the data[11], we have chosen to leave the high inflation years out of the sample. The second reason is that after the mid 1990s most of the countries of our sample had already liberalised capital markets, which presumably reduced the effect of capital controls and transaction costs on UIP.

We have chosen to use the Treasury Bill Rate for developed economies and also Brazil and Mexico. Deposit rates were used for Argentina, Chile and Turkey because of data availability. The annualised monthly interest rate was transformed into a compounded quarterly rate. Quarterly exchange rates changes were calculated, in percentage terms, using

[10] We also used Germany as the reference country, however, tests retrieved very similar results as to those using the US and we opted not to report them.

[11] During the 1980s and in the early 1990s, both inflation and interest rates show a very erratic behavior of *nid*s that for some observations take values beyond one billion. Observations for which *nid*s take implausible values correspond to a significant proportion of the sample for developing countries, especially Argentina and Brazil. Also, note that Argentina and Brazil carried out profound reforms at the beginning of the 1990s which were responsible for the monetary stabilisation.

data on end-of-period exchange rates. Output is measured as the natural logarithm of industrial production. The output gap was obtained by first de-trending the series using the Hodrick-Prescott filter. The value used for the smoothing coefficient is 126.400, which is recommended for monthly observations. We then subtracted the trend from the natural logarithm and multiplied by one hundred in order to obtain a numerical value comparable to the percentages of the other variables. Output changes were calculated as the monthly percentage changes in the index of industrial production. With regard to output, the data finish in 2000M5 for Argentina, 2004M1 for Brazil and Italy, 2004M2 for Chile, Mexico, Turkey, Spain and the UK and finally, 2003M10 for Germany.

3.2. Estimations of Equations (2) and (6)

We started by performing a sequence of unit root tests on the series of *nids* in order to investigate its univariate properties. Because of the low power of ADF test, we also checked for unit roots using the Kwiatkowski *et al.* (1992), Elliott *et al.* (1996), Perron (1997) and Elliott (1999) tests. We calculated the number of lags using a general to specific (joint F-test for the significance of the parameters at the 5% confidence level) approach starting from 12 lags. As can be seen in Tables 1 and 2, we found overall support for stationarity. Turkey and the UK were the only countries in which we could not reject the unit root or to accept stationarity. The low power of unit root tests, as mentioned before, and the relatively short period of our sample - for which a cointegrating relationship may not hold - can be driving the non-stationary results. Nonetheless, the findings that will be presented for these countries should be read with caution. Regarding Perron (1997) tests, the break date is endogenously chosen to be the one in which the t-statistic for testing the null hypothesis of a unit root is smallest among all possible break points. The breaks of emerging economies, presented in Table 2, seem to be related to the effect of domestic and financial crises. The introduction of

the Euro as a medium of exchange in 2002M1 is apparently reflected in the form of a date break of developed countries, with the exception of the UK.

Table 3 presents parameter estimates of equation (2) in which the authorities lean against exchange rates[12]. The findings are similar to Christensen (2000) who performs his tests for the US, Germany and Japan. The coefficient λ^S is close to zero in both studies (in most cases) and generally insignificant. On the other hand, σ is significant and lies between 0.96 and 0.98 in Christensen (2000) and among 0.75 and 1.00 in our tests. Results, *prima facie*, indicate that Central Banks of both developed and emerging economies practice interest rate smoothing but do not react to exchange rate changes. The above finding also points out to a very high degree of persistence in interest rates and for Germany the process seems to follow a unit root (which, as shown before, had been rejected).

In short, the results do not corroborate the intuition of McCallum (1994a), as he predicted $\lambda^s \cong 0.20$ and $\sigma \cong 0.80$. Regarding diagnostic tests, serial correlation was often eliminated through the introduction of additional lags without any considerable change in the parameters[13]. F-tests of autoregressive conditional heteroscedasticity (ARCH) with 4 lags show that variance is time dependent in most countries. Because of normality and conditional heteroscedasticity problems, we tested equation (2) using ARCH and generalised autoregressive conditional heteroscedasticity (GARCH) models. However, the findings above are roughly repeated.

Table 4 presents the results from the estimations of equation (6). The parameter λ^p, associated with price changes, is not significant in any developed country. For emerging markets, however, λ^p is significant and close to 0.2. Also, the degree of interest rate

[12] Notice that there is no identification problem since equation (2) excludes variable ξ_t, while equation (3) excludes the predetermined variable nid_{t-1}, hence, each equation excludes one variable that is present in the other equation. Variables ξ_t and nid_{t-1} give additional information about the variability of the dependent variable which allows the distinction between the estimation of the monetary policy reaction function and UIRP.

[13] For the sake of space and conciseness we did not to report these results.

smoothing diminishes when the policy reaction function is modified. Regarding residuals, serial correlation was eliminated by adding more lags to the regression without any important change in the parameter values. We also ran a battery of OLS regressions using a combination of exchange rates, output changes and output gap on the right-hand side of equation (2). The inclusion of more variables in the modified policy reaction function does not influence the coefficient λ^p to a great extent. None of these variables appeared to be systematically significant, thus, results are not shown.

Problems of normality and conditional heteroscedasticity were common. When the number of lags in the squared residuals is reduced to 1 or increased to 12, then conditional volatility seems to be a feature of nearly all estimated equations. We dealt with conditional heteroscedasticity and normality altogether by making use of ARCH and GARCH models and a generalised error distribution for the error term, which is appropriate when the time series is leptokurtic or exhibits tallness[14] [see Nelson (1991)]. We have chosen the type of the model [between ARCH, GARCH, EGARCH (exponential GARCH), LGARCH (leveraged GARCH) etc] according to the significance of the parameters on the conditional variance equation. The lags of the conditional variance were selected according to tests of ARCH effects. We started with a (p,q) model (1,1) in which p represents the order to the autoregressive term and q the number of lags of the moving average term. More lags were included if F-tests indicated serial correlation in the squared normalised residuals. The number of autoregressive terms in the mean equation was selected by analysing serial correlation in the normalised residuals. In the estimations that follow, the model for Argentina, Mexico and Turkey corresponds to an ARMA(1,1) while an AR(1) is used for Brazil and Chile. Finally, we added dummies for the largest outliers, which are likely associated with foreign or domestic financial crises, in order to ameliorate remaining normality problems.

[14] This distribution is represented by a function that captures the leptokurtic properties of a series, in other words, the density function has a higher probability for extreme events.

In Table 5, we report the coefficients of the mean equation using the model represented by (6) which was estimated taking conditional volatility into consideration. Results for developed countries are similar to the ones using OLS and are not reported. Comparing the λ^p coefficients of emerging economies in Table 4 with the ones in Table 5, it is possible to see that the degree of leaning against price changes diminishes when volatility clustering is taken into account. However, the estimates for Argentina, Brazil and Mexico are still relatively close to the presumption of 0.2.

The variance of the series significantly influences not only the squared error but also the mean equation in three countries. The ARCH-in-mean term ranges from a small positive value of 0.18 in Mexico to a higher one of 0.97 in Argentina. Hence, it could be said that monetary authorities consider uncertainty and/or risk[15], arising from the volatility of *nids*, when formulating their policy. A one unit increase in the variance of the *nids* of Argentina, for example, increases the difference between the quarterly nominal return of that country and the US by nearly 1%. Diagnostic tests show that there are no ARCH effects - at least in the four lags of squared normalised residuals – and the Ljung-Box Q-Test statistic reveals that there is no serial correlation in the normalised residuals. Jarque-Bera tests show that the distribution of the error term is normal[16].

Asymmetry in the squared residuals seems to be a common feature of most estimates. The fact that EGARCH or LGARCH models were the most relevant for all countries except Turkey shows that conditional variance depends asymmetrically on negative and positive shocks. The EGARCH model put forward by Nelson (1991) allows not only the magnitude

[15] Granger (2002) suggests that variance cannot be a measure of risk premium only, but also uncertainty: agents diversify their portfolio in order to avoid unexpected large losses but not large gains.

[16] With respect to the dummies, we used two in Argentina for the months 2001M9 and 2001M11, which may be reflecting the start of the financial crises that led to the free floating of the Peso; we also needed dummies for the months 1998M1, 1998M11 and 1998M12 in Brazil as they are likely mirroring the disproportionate responses of monetary policy to the threat that the Asian and Russian crises posed to the peg of the Brazilian Real; for Chile and Mexico we had to use a single dummy for 1998M11 which probably indicates a radical monetary response to the Russian crisis; finally, we had dummies for 1996M11, 2000M11 and 2001M12 in Turkey, the last two years possibly associated with the financial crisis that culminated with the free floating of the Lira.

of unanticipated returns to determine the variance but also the signs of the residuals. The leverage term of the LGARCH model is a dummy that captures asymmetry in the conditional variance. The dummy takes the value one when the residual is negative and zero otherwise.

We also performed several tests alternating and combining exchange rates, output changes and output gap in the policy reaction function, but results did not show any major difference in relation to the parameters reported in Table 5.

Changes in Monetary Policy

Because exchange rate regimes changed during the sample period, the "Lucas Critique" may apply in one sense: alterations in the way economic policy is conducted would influence the predictions of the models above. Hence, we decided to test whether changes in exchange rate systems have influenced the results previously found. We performed these tests using the following specification:

$$\begin{aligned} nid_t = {} & \alpha + \sigma_1 nid_{t-1} + D\sigma_2 nid_{t-1} + \lambda_1^p \Delta p_t + \\ & D\lambda_2^p \Delta p_t + \lambda_1^s \Delta s_t + D\lambda_2^s \Delta s_t + \theta_1 y_t^g + D\theta_2 y_t^g + \mu_t \end{aligned} \qquad (7)$$

where y_t^g is the output gap and μ_t is an error term with zero mean and variance following a heteroscedastic process; the dummy variable *D* takes the value zero in the first period and one in the second; the coefficients $\sigma_1, \sigma_2, \lambda_1^p, \lambda_2^p, \lambda_1^s, \lambda_2^s, \theta_1$ and θ_2 stand for the response of the monetary authorities to the variables that they are associated, for the corresponding periods 1 and 2 as shown in the subscripts. The first period corresponds to the Currency Board in Argentina (which finished in 2002M1) and managed exchange rates in Brazil (terminating in 1999M1) and also in Turkey (ending in 2001M2), while the second is the floating system. For Chile, the first period corresponds to the floating system with bands,

which was abandoned in 1999M9 for a flexible regime. Mexico had a floating system during the whole sample period and, for this reason, we did not select any date break for it.

Results are presented in Table 6. An EGARCH was the most relevant model for Brazil and Turkey, a LGARCH for Chile and Mexico and a GARCH model was the most adequate to Argentina. We used a generalised error distribution for all the emerging economies, except Brazil for which Jarque-Bera tests showed that residuals are normal[17]. Finally, F-tests rejected that ARCH effects or serial correlation remained in the normalised residuals up to lag four.

Tests including the output gap for Argentina are inappropriate because they entail a loss of more than half of the observations due to data availability. As can be seen in Table 6, interest rate smoothing decreases significantly in this country and reaction against inflation seems to be stronger during the floating period. For Brazil, all coefficients but the dummies of the interest rate smoothing variable and the output gap are significant. During the pegged system, an exchange rate change of 1% corresponded to an increase of 0.37% in the mean of *nids* (keeping other variables constant) which indicates a significant degree of leaning against the wind. Prices explain *nids* in the two periods but the sign of the parameter in the first is negative, which is puzzling. In the floating period, exchange rate changes do not have an impact on *nids* but inflation provokes a response of 0.12% in *nids*. *Nids* seem to react to the output gap during the managed float (they decrease as the gap increases) but are not significant during the floating period. Chile presents the smallest degree of interest rate smoothing which is around 0.50. Reaction to price changes, exchange rate changes and output gap is apparent only in the first period. After the structural break, there is interest rate smoothing but results suggest that there is no reaction to any other variable. Results for

[17] We had a dummy for Argentina in 2001M9 while for Chile and Mexico the dummy was for 1998M11; in Turkey we used one in 2001M2 in addition to another one for 1998M11. As can be seen in Table 4, the Jarque-Bera tests show that there could still be some concern for normality problems in Argentina, Mexico and Turkey. We opted for showing the present results because as more dummies are included and normality problems are eliminated, the sizes of estimated parameters do not change significantly but there is a loss in terms of degrees of freedom.

Mexico are robust with regard to the choice of the technique. Reaction against price changes is significant and in the order of 0.22. There is weak evidence of reaction to exchange rates or output for this country. According to the results, Turkey experienced a high degree of interest rate smoothing during the managed float with small reaction of *nids* to any of the variables (prices, exchange rates and output). The degree of interest rate smoothing diminishes during the floating period and there is a significant increase in the reaction against price changes. Finally, the variance of the series of *nids* positively influences the mean equation in Argentina and Chile but is not significant to explain the mean of Brazil, Turkey and Mexico. The ARCH-in-mean term is 0.18 and 0.32 in Argentina and Chile, respectively. In general, results suggest that shocks affect variance (which can be understood as a measure of uncertainty and/or risk) in a non-linear way and also that this variable is taken into account during the formulation of monetary policy.

The results seem to provide support for the idea that there is leaning against the parity for at least one period in all countries. The results are more solid for Argentina, Brazil, Mexico and Chile, the last country only in the first period. Brazil seems to lean against the wind during the managed float and react to inflation subsequently. There is evidence of strong reaction against price changes during the floating period in Argentina. According to the results, Chile leant against the parity by reacting to price changes until 1999M9. Evidence of monetary response in this country during the pure floating period is loose. Apparently, Mexico responded to inflation only. With the exception of Brazil most countries appear not to react to exchange rate changes, using both OLS and maximum likelihood conditional volatility estimations. A similar result (of nearly no policy reaction) is verified for the coefficients associated with the output gap and to output changes (the latter is not reported).

The importance of the output gap on the estimated policy reaction functions is negligible or the parameter is insignificant. One reason might be associated with the fact that price

stability is generally the Central Bank's main concern[18]. Another possible justification is that output gap is not directly observed and, hence, imperfect measurements are not able to capture the statistical relationship between the two variables. Similarly to output, exchange rate changes are generally not significant in the estimated policy reaction functions. The importance of the exchange rate channel to the formulation of monetary policy is directly related to the degree of pass-through. Devereux and Engel (2002, p.914), for example, explained that "low pass-through of exchange rates might imply high exchange rate volatility in equilibrium.". If the exchange rate has little effect on domestic expenditure, then it might take large changes in this variable to achieve equilibrium after some shock to fundamentals. Obstfeld (2004) presents a model showing that a free exchange rate system might be desirable, even if this variable has no expenditure-switching role, because it releases domestic authorities from using interest rates as a stabilisation tool. These ideas could explain why the statistical link between *nids* and exchange rates is so weak, especially during the floating period.

As a matter of fact, the fear of floating literature is incipient and more research needs to be carried out. We speculate whether Central Banks may only fear large swings in the exchange rate. In special, monetary authorities of emerging economies would drastically increase interest rates when there is a high probability of depreciation and reduce it slowly when the threat fades away. When pursuing its inflation target, the Central Banks of those countries would only be afraid of large upward swings on the exchange rate and welcome appreciations as it helps in pushing inflation down. Alternatively, there would not be any significant reaction to exchange rate appreciations if prices are rigid downwards.

3.3. Ex Post Deviations from UIP

[18] Theoretically, even if the Central Bank follows the strict rule of targeting price changes, it would still react to output to the extent at which this variable contains information about inflation [see Eichengreen (2002)].

The final step is to estimate the parameter ρ from $\xi_t = \rho\xi_{t-1} + \upsilon_t$. The variable ξ represents deviations from UIP, which can stem from frictions in financial markets, and is supposed to be autoregressive. For this reason, we apply unit root tests to investigate the dynamic properties of this variable. ADF statistics are presented in Table 7. We calculated the number of lags using a general to specific as in section 3.2. ADF statistics shows that the unit root is rejected at the 5% significance level in all cases, except Mexico. We also tested for unit roots in the *ex post* UIP deviations of Mexico using the Kwiatkowski *et al.* (1992) and Elliott *et al.* (1996) but there was no support for stationarity. Tests using the approach of Elliott (1999) provide the only statistic that is able to reject the null at the 10% confidence level.

The estimated root is significant and ranges from 0.50 to 0.70, which implies that the dynamic properties of UIP *ex post* deviations, often interpreted as risk premium [see Kugler (2000) and Berk and Knot (2001), for instance], influence the equilibrium *nid* by a substantial extent. A robust aspect of the results is that estimated parameters of the autoregressive term are significant and the speed of convergence is similar in magnitude. No assumptions were made for this coefficient in McCallum (1994a) and no estimation was performed in Christensen (2000), so it is not possible to compare the results.

4. Conclusion

The findings using the original model of McCallum (1994a) do not support his intuition as there is no evidence of leaning against the wind for most countries in most periods. The version of the policy reaction function with reaction against prices lends support to the idea of leaning against the parity using OLS estimation but only for emerging economies. The size and the sign of the parameters surprisingly resemble the values stated in McCallum (1994a). This model gives similar results when heteroscedasticity and normality problems are

taken into consideration – Turkey and Chile are the exceptions. Then, we tested a policy reaction function including exchange rate changes, output gap and we also checked whether parameter values changed by considering structural breaks (which were identified by changes in exchange rate regimes). We found evidence of leaning against the parity in all emerging countries in at least one period. Finally, evidence of interest rate smoothing was found to be robust for all countries and the dynamic properties of UIP *ex post* deviations, normally interpreted as risk premium, influenced the equilibrium *nid* of emerging economies to a reasonable extent. The impact of positive and negative shocks on the conditional variance is asymmetric. Furthermore, there is evidence that monetary authorities allow for uncertainty and/or risk in the policy reaction function.

The paper presented results that lend support to the view that policy actions simultaneously interact with the decisions of arbitrageurs and change the result predicted by international parity conditions, such as UIP and also, by consequence, the real interest rate parity hypothesis. McCallum (1994a) showed that smoothing interest rates and leaning against the wind under certain assumptions about parameter values imply that high nominal returns are associated with exchange rate appreciations, instead of (*ex post*) depreciations. Christensen (2000) did not find support for his model in a sample of developed countries. We have shown that monetary authorities play a role in the determination of *nids* in emerging countries if reaction against prices and exchange rate regime changes are taken into account into the estimation of the monetary policy reaction function.

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Table 1
Unit Root Tests on *Nids*

		ADF	KPSS	ERS	Elliot (1999)
	nº of lags	t-ratio	η_μ	DF-GLS	DF-GLS$_u$
Argentina	8	-2.32	0.43*	-2.27*	-2.33
Brazil	7	-3.55*	0.53*	-0.47	-2.78*
Chile	5	-1.83	1.31	-1.65**	-1.89
Mexico	6	-2.26	1.23	1.087	-1.17
Turkey	3	-1.80	1.33	-1.61	-1.88
France	12	-1.58	0.25*	-1.12	-1.51
Italy	12	-2.88**	0.36*	-1.03	-2.48**
Spain	8	-2.61**	0.25*	-0.99	-2.26
UK	1	-1.04	2.97	-0.47	-1.12
Germany	12	-1.74	0.66**	-1.80**	-1.88

Notes: * indicates rejection of the null of a unit root at the 5% confidence level for the ERS (1996) and Elliott (1999) tests and non-rejection of the null for the KPSS test at 95% confidence level. ** indicates rejection of the null of a unit root at the 10% confidence level for the ERS (1996) and Elliott (1999) tests and non-rejection of the null for the KPSS test at 99% confidence level.

Table 2
Perron (1997) Unit Root Tests on *Nids*

	Lags	Break Date	T-ratio
Argentina	12	2002:01	-5.98*
Brazil	7	1999:06	-3.93
Chile	3	1999:05	-6.10*
Mexico	1	1996:04	-5.37*
Turkey	10	2001:02	-3.72
France	0	2000:12	-4.22**
Italy	12	2001:01	-3.94**
Spain	8	2001:01	-4.38*
UK	3	1998:11	-1.48
Germany	12	2001:01	-6.43*

Notes: * indicates rejection of the null of a unit root at the 5% confidence level and ** indicates rejection of the null of a unit root at the 10% confidence level.

Table 3
Results obtained from the OLS estimation of Equation (2)

Developed Countries

	France	Italy	Spain	UK	Germany
Intercept	-0.004	-0.002	-0.002	0.013	0.005
	(0.009)	(0.008)	(0.005)	(0.008	(0.007)
σ	0.96*	0.96*	0.97*	0.97*	1.00*
	(0.023)	(0.014)	(0.011)	(0.023)	(0.019)
λ^s	-0.003	0.000	-0.002	-0.002	0.001
	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
R^2	0.96	0.98	0.98	0.95	0.97
Serial Correlation[a]	18.80	17.10	28.58	18.83	14.77
	[0.93]	[0.14]	[0.00]	[0.09]	[0.25]
Functional Form[b]	0.92	0.13	0.28	0.00	0.18
	[0.34]	[0.71]	[0.59]	[0.93]	[0.66]
Normality[c]	50.56	0.95	6.88	2.50	5.90
	[0.00]	[0.62]	[0.03]	[0.28]	[0.05]
Heteroscedasticity[d]	1.91	8.34	0.03	1.05	2.52
	[0.17]	[0.00]	[0.85]	[0.31]	[0.11]
F-ARCH (4 lags)	0.30	5.34	2.95	4.93	0.97
	[0.87]	[0.00]	[0.02]	[0.00]	[0.42]

Emerging Markets

	Argentina	Brazil	Chile	Mexico	Turkey
Intercept	0.15	0.91*	0.31*	0.29*	2.96*
	(0.171)	(0.323)	(0.108)	(0.149)	(0.92)
σ	0.92*	0.83*	0.75*	0.89*	0.76*
	(0.038)	(0.051)	(0.066)	(0.027)	(0.059)
λ^s	-0.006	-0.008	0.002	0.002	0.079*
	(0.006)	(0.011)	(0.014)	(0.018)	(0.024)
R^2	0.85	0.71	0.56	0.91	0.74
Serial Correlation[a]	47.33	21.23	16.32	13.52	23.20
	[0.00]	[0.05]	[0.18]	[0.33]	[0.03]
Functional Form[b]	0.05	7.24	2.19	3.29	19.58
	[0.82]	[0.00]	[0.14]	[0.07]	[0.00]
Normality[c]	773.85	3910.6	22.52	695.41	943.98
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Heteroscedasticity[d]	9.27	8.54	12.11	14.14	6.65
	[0.00]	[0.00]	[0.00]	[0.00]	[0.01]
F-ARCH (4 lags)	12.21	0.10	1.63	0.17	5.37
	[0.00]	[0.98]	[0.17]	[0.96]	[0.00]

Notes: * denotes significance at 5% level, standard errors are in parenthesis and significance levels in brackets. With regard to diagnostic tests: a) Lagrange multiplier test of residual serial correlation for 12 lags; b) Ramsey's RESET test using the square of the fitted values; c) Based on a test of skewness and kurtosis of residuals; d) Based on the regression of squared residuals on squared fitted values.

Table 4
Results obtained from the OLS estimation of Equation (6)

Developed Countries

	France	Italy	Spain	UK	Germany
Intercept	-0.007	-0.026	-0.009	0.005	-0.002
	(0.012)	(0.018)	(0.008)	(0.009)	(0.007)
σ	0.97*	0.95*	0.97*	0.99*	0.99*
	(0.022)	(0.014)	(0.011)	(0.002)	(0.017)
λ^p	0.005	0.042	0.009	0.009	0.018
	(0.022)	(0.028)	(0.009)	(0.008)	(0.011)
R^2	0.95	0.98	0.98	0.95	0.97
Serial Correlation[a]	17.21	15.26	28.99	19.63	14.06
	[0.14]	[0.23]	[0.00]	[0.07]	[0.29]
Functional Form[b]	1.19	2.08	0.01	0.08	0.48
	[0.27]	[0.15]	[0.92]	[0.77]	[0.49]
Normality[c]	96.91	1.64	5.86	2.28	6.98
	[0.00]	[0.44]	[0.05]	[0.32]	[0.03]
Heteroscedasticity[d]	1.67	10.73	0.13	0.47	2.11
	[0.19]	[0.00]	[0.72]	[0.49]	[0.15]
F-ARCH (4 lags)	0.16	4.00	1.63	4.58	0.87
	[0.95]	[0.00]	[0.17]	[0.00]	[0.48]

Emerging Markets

	Argentina	Brazil	Chile	Mexico	Turkey
Intercept	0.076	0.77*	0.11	0.09	2.45*
	(0.015)	(0.319)	(0.119)	(0.118)	(0.863)
σ	0.85*	0.76*	0.70*	0.57*	0.66*
	(0.035)	(0.099)	(0.064)	(0.045)	(0.067)
λ^p	0.20*	0.22*	0.25*	0.46*	0.25*
	(0.037)	(0.058)	(0.079)	(0.058)	(0.059)
R^2	0.88	0.73	0.60	0.94	0.76
Serial Correlation[a]	47.40	8.01	18.25	14.96	21.29
	[0.00]	[0.78]	[0.11]	[0.24]	[0.05]
Functional Form[b]	2.43	6.08	0.00	0.07	3.72
	[0.12]	[0.01]	[0.94]	[0.79]	[0.05]
Normality[c]	270.25	2967.6	27.87	760.12	1331.3
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Heteroscedasticity[d]	15.69	11.99	9.45	2.30	0.01
	[0.00]	[0.00]	[0.00]	[0.13]	[0.93]
F-ARCH (4 lags)	13.59	0.21	1.18	0.42	4.45
	[0.00]	[0.93]	[0.32]	[0.99]	[0.00]

Notes: * denotes significance at 5% level, standard errors are in parenthesis and significance levels in brackets. With regard to diagnostic tests: a) Lagrange multiplier test of residual serial correlation for 12 lags; b) Ramsey's RESET test using the square of the fitted values; c) Based on a test of skewness and kurtosis of residuals; d) Based on the regression of squared residuals on squared fitted values.

Table 5
Conditional Heteroscedasticity Estimation of Equation (6)

	Argentina	Brazil	Chile	Mexico	Turkey
Model	LGARCH(1,1)	LGARCH(1,1)	EGARCH(1,1)	EGARCH(2,2)	ARCH(1)
Intercept	0.03 (0.046)	0.82* (0.197)	0.12* (0.029)	0.11 (0.126)	-0.41* (0.074)
σ	0.59* (0.059)	0.78* (0.023)	0.64* (0.071)	0.76* (0.038)	0.99* (0.004)
λ^p	0.15* (0.028)	0.10* (0.039)	-0.02* (0.008)	0.17* (0.040)	0.03* (0.001)
Arch-in-Mean	0.97* (0.275)	-0.23 (0.488)	0.62* (0.184)	0.18* (0.475)	--
Jarque-Bera	3.996 [0.135]	0.410 [0.814]	1.717 [0.423]	2.675 [0.262]	2.918 [0.232]
F-ARCH (4 lags)	0.640 [0.635]	0.547 [0.701]	1.613 [0.177]	0.306 [0.835]	0.706 [0.589]
Ljung-Box Q-Test (4 lags)	1.823 [0.609]	1.248 [0.741]	5.152 [0.161]	3.9153 [0.271]	2.358 [0.501]

Note: * denotes significance at 5% level, standard errors are in parenthesis and significance levels are in brackets.

Table 6
Conditional Heteroscedasticity Estimation of Equation (7)

	Argentina	Brazil	Chile	Mexico	Turkey
Model	GARCH(1,1)	EGARCH(1,1)	LGARCH(1,1)	LGARCH(1,1)	EGARCH(1,1)
α	0.00	0.89*	0.12*	0.29*	2.07*
	(0.012)	(0.151)	(0.047)	(0.104)	(0.401)
σ_1	0.85*	0.75*	0.52*	0.71*	0.87*
	(0.039)	(0.055)	(0.063)	(0.038)	(0.020)
σ_2	-0.19*	-0.01	0.13	--	-0.11*
	(0.039)	(0.07)	(0.11)		(0.022)
λ_1^P	0.08*	-0.14*	0.37*	0.22*	0.00
	(0.028)	(0.085)	(0.069)	(0.044)	(0.018)
λ_2^P	0.19*	0.26*	-0.39*	--	0.09*
	(0.071)	(0.092)	(0.071)		(0.029)
λ_1^s	--	0.37*	0.05*	-0.02*	0.04*
		(0.174)	(0.025)	(0.008)	(0.014)
λ_2^s	0.00	-0.36*	-0.05*	--	-0.03*
	(0.009)	(0.175)	(0.026)		(0.016)
θ_1	--	-0.10*	0.04*	0.001	-0.02*
		(0.037)	(0.015)	(0.006)	(0.009)
θ_2	--	3.73	-0.04*	--	-0.04*
		(4.54)	(0.016)		(0.015)
ARCH-in-Mean	0.18*	-0.10	0.32*	0.00	-0.10
	(0.073)	(0.227)	(0.177)	(0.151)	(0.119)
Jarque-Bera	7.901	0.712	4.52	4.74	4.890
	[0.019]	[0.700]	[0.104]	[0.093]	[0.087]
F-ARCH (4 lags)	0.79	0.41	0.39	1.77	0.43
	[0.53]	[0.801]	[0.811]	[0.142]	[0.785]
Ljung-Box Q-Test	5.45	2.09	4.41	2.32	2.74
(4 lags)	[0.141]	[0.554]	[0.220]	[0.507]	[0.433]

Note: * denotes significance at 5% level, standard errors are in parenthesis and significance levels in brackets.

Table 7
Unit Root (ADF) tests on ξ_t

	Argentina	Brazil	Chile	Mexico	Turkey
Nº of lags	4	4	4	7	1
ADF statistic	-4.437*	-4.477*	-4.704*	-2.610	-5.871*
ρ (estimated root)	0.67*	0.50*	0.60*	0.70*	0.58*

Note: * represents rejection of the null at 5% level for the ADF statistic, and significance at 5% for the estimated root.