Are real interest differentials caused by frictions in goods or assets markets, real or nominal shocks?

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Abstract

The variance of real interest rate differentials (*rids*) is decomposed between *ex post* deviations from relative purchasing power parity and uncovered interest rate parity (UIRP) for a set of emerging markets from 1995M5 to 2004M3. The results point out to nominal interest rate differentials and *ex post* deviations from UIRP as the main source of volatility in *rids*. In order to uncover the dynamic effects of real and monetary disturbances, I estimated a bivariate VAR with *rids* and nominal interest rate differentials. Forecast error variance decomposition using short run restrictions on the VAR strongly supports the claim that money shocks are unable to explain the variability of *rids* at longer horizons. Long-run restrictions results in real shocks as the likely cause of *rids*. Analysis of impulse response functions demonstrates that the net impact of a (one standard deviation) real shock on *rids* after 36 months is large.

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1. Introduction

Uncovered Interest Rate Parity (UIRP) with rational expectations and relative Purchasing Power Parity (PPP^r) entail the Real Interest Rate Parity Hypothesis (RIPH) [Roll (1979)]. The conclusion regarding the existence of *ex post* real interest rate differentials (*rids*, hereafter) across countries since the seminal papers of Mishkin (1984) and Cumby and Obstfeld (1984) is not decisive [see for example Obstfeld and Taylor (2002)]. The usual finding is that *rids* are autoregressive and relatively short-lived. The aim of the current paper is to investigate the general causes of *rids*. For this purpose, I use a selected sample of emerging markets in which latest evidence indicated that *rids* (in relation to the US) mean-revert to a positive equilibrium and have asymmetric behaviour [see Ferreira and Leon-Ledesma (2003)].

Departures from RIPH can be explained by *ex post* deviations from PPP^r and UIRP. Hence, a question that arises is whether *rids* are caused by frictions in goods or assets markets? Another interrelated question is if real shocks (changes in risk perception or productivity increases, for example) are more important than nominal shocks (such as unexpected changes in money supply, for instance) to explain deviations from interest parity. These questions are relevant because RIPH is based on the existence of frictionless markets and *rids* reflect the degree of market integration. The answers might be of practical importance for researchers as well as for policy makers. For example, stabilising the variance of *rids* can be a target of monetary policy in itself[1]. If *rids* are very volatile, returns are unstable and investors dislike variance. The higher the variance, the smaller is the incentive to invest in a bond and the greater must be its return. Hence, policy makers may want to offset shocks that cause great variability. Also, high *rids* can impose heavy costs to an economy - because of interest payments on the public, domestic and foreign debt - so

[1] See Iwata and Tanner (2003) for evidence on the trade-off between exchange rate and interest rate volatility in developing countries.

unveiling the causes and understanding their dynamics is essential to design the appropriate macroeconomic policies to change differentials.

There are also theoretical issues motivating the work. Variance decompositions can shed light on the nature of the relationship between *rids* and real exchange rates. There has been a debate on whether this relationship holds since Frankel (1979). Evidence can be non-supportive as Meese and Rogoff (1988), Edison and Pauls (1993), MacDonald (1997), Breedon et al. (1999) and Isaac and de Mel (2001) or favourable as Astley and Garrat (2000), Chortareas and Driver (2001), Macdonald and Nagayasu (2000) and Jin (2003). Because of Balassa-Samuelson effects, the sign of an impact of a real shock on exchange rates (and *rids*, as I will explain) is undetermined and depends on the type of the disturbance and the sector of the economy that is hit. The proposed tests can help to clarify this issue because, as MacDonald and Ricci (2003) observed: "*real interest rate differentials may also reflect productivity differentials: to the extent that the measure employed to proxy for the Balassa-Samuelson effect is not perfect, the real interest differential may help capture this empirically*." (Pps. 4 and 5, emphasis added).

I focus on the importance of the international parity conditions on the determination of *rids*. The broad question is whether *rids* can be explained by *ex post* deviations from PPP^r and UIRP and to which extent. The main objective is to separate out the driving sources of volatility in the variance of *rids*. The second goal of the paper is to characterise the dynamic response of *rids* to real and nominal disturbances and to breakdown its variability according to these two types of shocks.

The paper presents further evidence on a higher degree of friction in assets rather than goods' markets and the predominance of real shocks in the path of *rids* for a set of emerging economies. To my knowledge, no work has performed innovation accounting on *rids*, hence the tests are innovative in this sense. The work also complements papers on the relationship of real exchange rates and *rids* by reinforcing the finding of no correlation between variables. The rest of the paper is organised as follows. Section 2 describes the methodology involved in the tests and discusses the identifying restrictions; Section 3 explains the data and presents the results. Section 4 concludes.

2. Methodology and Theory

The first method is based on Cheung et al. (2003) who separated the variance of *rids* between deviations from PPP^r and UIRP using the relationships given by RIPH

$$rid_t = (i_t - i_t^* - \Delta s_t^e) - (\pi_t - \pi_t^* - \Delta s_t^e) \tag{1}$$

Where i is the domestic nominal interest rate and i^* is the foreign interest rate that matures at time t. The nominal exchange rate, S, is the domestic price of the foreign currency; the expected rate of depreciation is $\Delta s_t^e = \frac{S_t^e}{S_{t-1}} - 1$, with the superscript e denoting expected values and the subscript t standing for time. Domestic and foreign inflation rates are π_t and π_t^* respectively. One can decompose the variance of *rids* in the following ways

$$\begin{aligned} Var(rid_t) &= Var(i_t - i_t^* - \Delta s_t) + Var(\pi_t - \pi_t^* - \Delta s_t) - 2Cov(i_t - i_t^* - \Delta s_t; \pi_t - \pi_t^* - \Delta s_t) \\ Var(rid_t) &= Var(i_t - i_t^*) + Var(\pi_t - \pi_t^*) - 2Cov(i_t - i_t^*; \pi_t - \pi_t^*) \end{aligned} \tag{2}$$

The second method consists in recovering the relevant parameters for innovation accounting using short and long-run restrictions on a bivariate VAR system of equations. UIRP imply a relationship between exchange rates and interest rates which allows one to categorise real and nominal factors as being the main sources of disturbances affecting *rids* and *nids.* This categorisation is based on the literature that applied variance decomposition to exchange rates [Rogers (1999), Enders and Lee (1997) and Astley and Garratt (2000), for example]. Ignoring intercept terms for simplicity:

$$rid_t = \varepsilon r_t + \varepsilon n_t \quad (3)$$

$$nids_t = \varepsilon r_t + \varepsilon n_t \quad (4)$$

Where real and nominal shocks are represented by $\varepsilon r_t, \varepsilon n_t$ respectively; disturbances are assumed to be iid $\mathrm{N}(0, \sigma_\varepsilon^2)$ in which σ_ε^2 represents variance. Sequences (3) and (4) can be represented as moving average processes

$$rid_t = \sum_{k=0}^{\infty} c_{11}(k) \varepsilon r_{t-k} + \sum_{k=0}^{\infty} c_{12}(k) \varepsilon n_{t-k} \quad (5)$$

$$nid_t = \sum_{k=0}^{\infty} c_{21}(k) \varepsilon r_{t-k} + \sum_{k=0}^{\infty} c_{22}(k) \varepsilon n_{t-k} \quad (6)$$

The letter *c* stands for the coefficients associated with the responses of *rids* and *nids* to shocks at each period *k*. The VAR representation is

$$\begin{pmatrix} rid_t \\ nid_t \end{pmatrix} = \begin{pmatrix} A_{11}(L) & A_{12}(L) \\ A_{21}(L) & A_{22}(L) \end{pmatrix} \begin{pmatrix} rid_{t-1} \\ nid_{t-1} \end{pmatrix} + \begin{pmatrix} e_{1t} \\ e_{2t} \end{pmatrix} \quad (7)$$

where e stands for the error terms, which are composite of the pure innovations $\varepsilon r_t, \varepsilon n_t$. UIRP with rational expectations minus expected inflation differentials gives

$$rid_t = \Delta q_t^e \tag{8}$$

where the expected change of real exchange rate depreciation is Δq_t^e. Considering perfect foresight in (8), the system in (7) can be rewritten using exchange rate changes in the place of interest rate differentials. This would be the standard set up of the tests of the literature that applies innovation accounting to exchange rates.

The Choleski decomposition imposes a contemporaneous restriction in (3) or (4) in order to recover the parameters of (5) and (6) from the estimates of the system in (7). The assumption is that a real shock does not have a contemporaneous impact on *nids*, a conjecture that is valid provided that real shocks affect prices instantaneously while interest rates are impacted after one lag[2]. Another interpretation is that policy makers react to a real shock after having more knowledge of its nature. The time elapsed for the reaction to take place is one month[3].

For the Blanchard and Quah (1989) decomposition I considered that the sum of nominal shocks has a zero impact on the series of *rids*

$$\sum_{k=0}^{\infty} c_{12}(k)\varepsilon n_{t-k} = 0 \tag{9}$$

this is explained in an appendix to this paper.

[2] I discarded the possibility that a nominal shock does not contemporaneously affect *rids* because it is logically inconsistent. The reason is that a nominal shock would have to impact interest rates and prices both at the same time and by the same magnitude, leaving *rids* at time *t* absolutely unchanged. The inconsistency arises because even if there is no initial impact on *rids*, there would be lagged effects.

[3] Monetary Policy Committee meetings in Brazil, for example, are realised monthly and, in several cases, interest rates cannot change until the day of the meeting.

Either identifying restriction (long-run or contemporaneous) depends on a set of assumptions that might not be entirely accepted. It is often attributed to the VAR literature, the use of incredible restrictions (assumptions) for identification. Nonetheless, as pointed out by Sims (1980), Faust (1998) and Faust et al. (2003), even incredible restrictions can result in useful analysis provided that reasonable economic interpretations can be given to the findings. Faust (1998), for example, has elaborated a way of checking for robustness of contentious restrictions by taking a particular assumption and checking "…all possible identifications of the VAR for the one that is the worst case for the claim, subject to the restriction that the implied economic structure produce *reasonable* responses to policy shocks." (pp. 209 - emphasis from the author). Then, he adds, "If in the worst case the variance share is small, then the claim is supported. If the share is large, then either the identifying information – the characterization of a *reasonable* policy shock – must be sharpened or we must view the issue as unsettled." (pp. 210). I performed and compared variance decompositions of *rids* using both short and long-run restrictions as a way to verify the "robustness" of the assumptions.

3. Results

The emerging markets of the sample comprise the small open-economies of Argentina, Brazil, Chile, Mexico and Turkey. I used the USA as the reference large economy for the calculation of the *rid*. The period of the tests corresponds to the interval that spans from 1995M5 to 2004M3.

The sample period starts in the mid 90s because harmonised data for the construction of *rids* for some countries did not exist before this period and also because after the mid-90s most of the countries had liberalised capital markets and had advanced substantially in their trade liberalisation process. Data on interest rates and end-of-period exchange rates was obtained from IMF's International Financial Statistics (IFS). I have chosen the Treasury Bill Rates for Brazil, Mexico, and deposit rates for Argentina, Chile and Turkey because data availability. The inflation rate is the rate of growth of the Consumer Price Index (CPI). I transformed the annualised monthly interest rate and the inflation rate into compounded quarterly rates and then subtracted the latter from the former. Quarterly exchange rates changes were calculated using data on end-of period exchange rates.

Table 1 presents some descriptive statistics of the differentials. Note that *rids* are smaller than *nids* in all countries with the exception of Argentina. The reason is that Argentina experienced deflation in many months. The highest differentials are in Turkey followed by Brazil, Mexico, Argentina and Chile.

Table 1
Some Descriptive Statistics of *Rids* and *Nids*

	Variable	Mean	Min	Max
	nids	2.23	0.24	19.99
Argentina	*rids*	1.74	-15.36	13.58
	nids	5.58	2.60	21.20
Brazil	*rids*	4.01	-2.24	14.39
	nids	1.31	-0.41	5.13
Chile	*rids*	0.87	-1.15	4.08
	nids	4.12	0.87	18.4
Mexico	*Rids*	1.27	-1.72	4.83
	nids	16.59	6.47	32.04
Turkey	*rids*	5.09	-2.8	21.33

In order to find out the order of integration of *nids* before running the VAR, I tested for the presence of unit roots using ADF, Kwiatkowski et al (1992) (KPSS), Elliot et al (1996) (ERS), Elliott (1999) and Perron (1997) tests. I found the optimal augmentation lags using a general-to-specific sequential criteria. I report t-ratios without a time trend because it was found to be insignificant in most cases.

Table 2
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

* 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 the 1% confidence level.

ADF test rejected the null of a unit root for Brazil; ERS test rejected for Argentina and Chile and Elliot (1999) only for Brazil. KPSS did not reject the null of stationarity for Argentina and also Brazil.

Graphical analysis in Figure 1 and the cumulative sum of recursive residuals suggest the existence of breaks in the series. I performed Perron (1997) tests using a model in which the series contain an innovational outlier with a change in the intercept[4]. This model can be represented as:

$$rid_t = a_0 + \theta DU_t + \lambda D(T_b)_t + a_1 rid_{t-1} + \sum_{i=1}^{p} \beta_i \Delta rid_{t-i} + \varepsilon_t, \qquad (20)$$

where T_b denotes the break date; $DU_t = 1(t > T_b)$ and $D(T_b)_t = 1(t = T_b + 1)$. The test is performed using the t-statistic for the null hypothesis that $a_1 = 1$. The results are reported in Table 3.

[4] Visual inspection of the data points out to only one break. The method that was used chooses the minimum of the t-statistic for testing the hypothesis that the parameter associated with the first lag of the autoregressive variable (in level) is equal to unit. Other methods retrieved analogous results.

Table 3
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

* indicates rejection of the null of a unit root at the 5%.

The date breaks retrieved by the tests seem to reflect the effects of domestic crises in Argentina, Brazil, Mexico and maybe in Turkey. The exogenous event most closely associated with the data break of Chile is the Brazilian financial crisis. In short, the unit root was rejected for Argentina, Chile and Mexico. Hence, there is evidence of stationarity in *nids* of all countries with the exception of Turkey. For this reason, I run a VAR for Turkey using *nids* in both levels and first difference.

Regarding variance decomposition, the results demonstrate that the share of *ex post* deviations from UIRP in the variance of *rids* is higher than the share of *ex post* deviations from PPP^r for Argentina, Brazil, Chile and Mexico.

Table 4
Variance Decomposition of *rids* between UIRP and PPP deviations

	Argentina	Brazil	Chile	Mexico	Turkey
Variance of:					
Rids	13.2	5.4	0.9	1.9	15.0
Deviations from UIRP	648.9	178.4	22.3	35.2	98.7
Deviations from PPP^r	541.5	175.8	21.4	29.8	105.7
% of *Rids*' variance:					
Deviations from UIRP	4930.9	3289.8	2603.8	1895.0	656.07
Deviations from PPP^r	4115.0	3241.3	2492.5	1601.0	702.7
-2cov(UIRP,PPP^r)	-8945.9	-6431.1	-4996.3	-3396.0	-1258.8

The high volatility of exchange rates is responsible for most part of the variance of individual parity conditions. A clear picture on the causes of deviations from RIPH emerges when *rids* are decomposed between *nids* and inflation differentials, as in Table 5. It becomes apparent that *nids* are the predominant source of variability for most *rids* of the sample. Inflation differentials account for a higher share of *rids*' variance only in Turkey.

Table 5
Variance Decomposition of *rids* between *nids* and inflation differentials deviations

	Argentina	Brazil	Chile	Mexico	Turkey
Variance of:					
Rids	13.2	5.4	0.9	1.9	15.0
Nids	14.0	7.7	1.0	11.3	25.4
Inflation differential	12.8	3.0	0.6	9.1	31.7
% of *Rids*' variance					
Nids	**106.6**	**142.4**	**117.3**	**606.1**	**168.7**
Inflation differential	**97.5**	**54.7**	**71.4**	**488.7**	**210.7**
-2cov(*Nids*, Inf. Differential)	-104.1	-97.1	-88.7	-994.9	-279.4

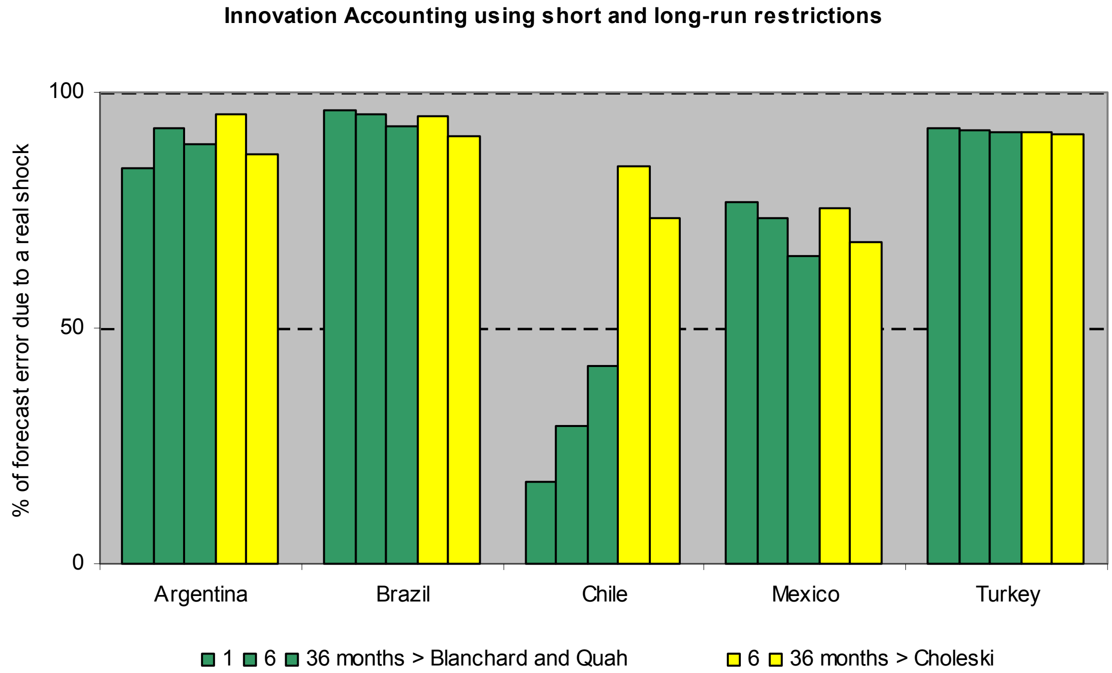
The covariance between *nids* and inflation differentials and the value of the correlations (the latter is not reported) indicate that the two variables have some degree of dependence. Interestingly, there is a lack of correlation between both *nids* and inflation differentials with respect to exchange rate changes (the exceptions are Argentina and Turkey, the latter in a smaller degree). If *nids* and inflation differentials do evolve according to processes that are independent from exchange rate variations, as shown by covariances and correlations, then there cannot be a significant relationship between *rids* and real exchange rates. This deduction is supported by the "no association" result of the empirical papers that analysed the link between the two variables.

In conclusion, the volatility of *nids* explains the majority of *rids*' variance in most economies. *Nids* seem to be fairly independent from exchange rate variations which point out to other factors explaining its behaviour and, by consequence, the dynamics of *rids*. Risk premium or the influences of monetary policy on deviations from UIRP, as pointed out by

McCallum (1994), are possible explanations[5]. Inflation differentials play a smaller role in explaining the variance of *rids* in emerging economies, with the exception of Turkey.

I turn to the findings of innovation accounting by first analysing forecast error variance decompositions[6].

Figure 2. Forecast Error Variance Decomposition of *Rids*



Note: The forecast error variance decomposition is the percentage of the mean squared error due to a real shock.

Figure 2 shows the percentage of variance explained by real shocks for some selected time-horizons: 1, 6 and 36 months for Blanchard and Quah (1989) and 6 and 36 for Choleski decomposition. Real shocks are the main source of variation in *rids* for all countries at all

[5] For instance, Calvo and Reinhart (2002) explained that some countries suffer from the "Fear of Floating". The story is that Central Banks of frightened economies would be scared of exchange rate volatility and would put too much weight on exchange rate stabilisation when setting interest rates. The authors have shown that the variance of nominal and real interest rates is high in emerging economies that experience low levels of inflation (estimated in about four times that of developed economies)

[6] I do not present and discuss the results of the VAR estimates as the primarily objective of the paper is to analyse forecast error variance decomposition and impulse responses. The optimal lag length was selected by a general to specific method using a likelihood ratio test for the exclusion of the last lag in each VAR equation, starting with 12. The lags chosen were Argentina (12), Brazil (10), Chile (10), Turkey (5) and (9) for the *nids* of Turkey in first difference. The tests were performed using the software RATS and the program var.src written by Norman Morin and available at Estima home-page http://www.estima.com/. Results are available with the author upon request.

horizons according to the Choleski decomposition. Blanchard and Quah (1989) reveals that, with the exception of Chile, the highest share of total variation in *rids* derives from a real shock.

Figure 3 presents impulse responses obtained through the use of Blanchard and Quah (1989) technique as short-run responses would be somewhat influenced by the contemporaneous restriction. Long run restrictions leave the short run dynamics of the VAR unconstrained or data-determined and structural theoretical explanations for variance decompositions and impulse responses can be made, as Clarida and Gali (1994) and Astley and Garratt (2000) emphasised.

It is important to note that a positive shock to the *rid* means that the expected exchange rate depreciation is higher than the one actually observed. It follows that the exchange rate depreciates by more than expected when there are no Balassa-Samuelson effects and the economy is subjected to an unexpected productivity increase (a positive real shock), hence *rids* diminish. On the other hand, *rids* increase if there are Balassa-Samuelson effects. The reason is that an unexpected productivity rise generates an unexpected appreciation. The channel by which risk affects *rids* is direct. Hence, an unanticipated increase in risk raise *rids*. Finally, a real demand shock leads to a permanent real appreciation and also enlarge *rids*.

Responses were normalised so each structural shock correspond to one standard deviation. As can be seen in Figure 3, a real shock causes a positive impact in both *rids* and *nids* of Argentina. The response of *rids* to nominal disturbances go to zero very quickly in Brazil but real shocks trigger a more persistent effect. The initial (and accumulated) effect of a real shock to both *rids* and *nids* is positive. After 3 years, a real shock adds 6.52 units to the sum of *rids* of Brazil. Impulse responses of Chile show that the first impact of a real shock is positive for *nids* but not for *rids*. The accumulated effect of a real shock after 36 periods is a rise of 1.94 units in the *rids* of Chile. A real shock originally increases *rids* and *nids* of Mexico. On the other hand, the initial effect of a nominal shock is ambivalent. After 36 months the accumulated impact of a nominal shock to *rid* dies out while a real shock effect sums up to 1.85 units. A positive shock (real or nominal) increases *nids* and *rids* of Turkey in the short run. After 3 years, a real shock increases *rids* by 5.71 units. Impulse responses of Turkey using *nids* in first difference provide a similar result.

The final impact of a real shock is considerably larger in Argentina (4.8 units)[7], Brazil and Turkey and slightly higher in Chile and Mexico. The reason for a higher accumulated impact than the initial increase might be related to frictions in financial markets or to the breakdown of rational expectations. Finally, while the sign of the accumulated impact of real shocks on *nids* is ambivalent, they are positive for *rids* of all countries with the exception of Chile. As the 1990's was a period characterised for productivity increases, this result, *prima facie*, lends support for Balassa-Samuelson effects[8]. Finally, nominal shocks can have different sorts of impacts on *rids* and *nids* in the short-run.

4. Concluding Remarks

Deviations from international parity conditions do not provide a clear picture on the causes of *rids* because exchange rate changes are very volatile and, in fact, cancel out in the composition of *rids*. The variance of *nids* explains most part of the volatility of *rids* for all countries, except Turkey. Recall that *rids* are calculated *ex post* so the aforementioned variance decomposition does not require any statistical test based on probabilities because *rids* are equal to *nids* subtracted from inflation differentials by definition. *Nids* seem to be fairly independent from exchange rate variations which signal to other factors explaining its

[7] It must be stressed, however, that the presence of outliers casts some doubt on the results for Argentina.
[8] See Lee and Tang (2003) for latest survey and evidence on the relationship between productivity and real exchange rates.

behaviour and, by consequence, the dynamics of *rids*. Risk premium and monetary policy behaviour that introduces persistence in *nids* [as in McCallum (1994)] are potential candidates. There seems to be no or small correlation between inflation differentials and exchange rate changes as well. These results may not be surprising as many empirical works did not find a significant relationship between interest rate differentials and exchange rates.

I found evidence of stationarity for all *nids* in the sample, with the exception of Turkey. Forecast error variance decomposition shows that real shocks explain most part of the variation in *rids* and the results are robust to either form of identifying restriction. The effect of a real shock tends to be amplified in the long run, reflecting the fact that, whenever differentials of developing economies start to grow, the tendency is for them to accumulate by more than the initial increase. This reinforces the findings of frictions in assets markets. The sign of the impact of real shocks on *nids* is ambivalent, but they are positive for *rids* of all countries with the exception of Chile. At the extent to which real shocks reflect productivity changes, this result provides support for Balassa-Samuelson effects. However, it must be stressed that the 1990's was also a period of various financial crises and the results of endogenous date breaks seem to reflect this fact. Finally, nominal shocks can have different sorts of effects on *rids* and *nids* in the short-run.

Arbitrage is supposed to be largely enforced by increased market integration. As the sample period follows the trade and financial liberalisation, one would expect that departures from parity conditions played a minor role in the composition of *rids*. This possibility is weakened if imperfect asset substitutability is a plausible conjecture for the financial markets. The findings of the present paper reveal the predominance of *nids* and real shocks in the path of *rids* for most countries which points out to deviations from UIRP as their driving source. The remaining puzzle is that *nids* seem to have no correlation with exchange rate changes.

Appendix

I start by showing the objective of the argumentation, for which I rewrite equation (5) from Section 2, considering, for simplicity, that the parameters of the moving average representation $c_{2,1}(k)$ $c_{2,2}(k)$ are all equal to 1. The real interest rate differential (*rid*) is then defined as

$$rid_t = \sum_{k=0}^{\infty} \varepsilon r_{t-k} + \sum_{k=0}^{\infty} \varepsilon n_{t-k} \tag{A1}$$

and I want to show that $\sum_{k=0}^{\infty} \varepsilon r_{t-k}$ is different from zero. As *rids* are stationary the difficulty lies not only in showing that $\sum_{k=0}^{\infty} \varepsilon n_{t-k} = 0$ but also that $\sum_{k=0}^{\infty} \varepsilon r_{t-k} \neq 0$.

First suppose that $\sum_{k=0}^{\infty} \varepsilon n_{t-k}$ is zero and rid_t is not zero. Trivial arithmetic is enough to conclude that $\sum_{k=0}^{\infty} \varepsilon r_{t-k}$ is different from zero. Thus, I need to justify this argument.

1) Assuming long-run money neutrality, $\sum_{k=0}^{\infty} \varepsilon n_{t-k} = 0$

It follows from Beveridge and Nelson (1981) univariate decomposition of real exchange rates, as performed in Clarida and Gali (1994), that

$$\hat{q} = q_{t-T} + E_{t-T} \sum_{k=0}^{T} \Delta q_{t-k} \tag{A2}$$

where the permanent element $\hat{q}$ is the expectation of q_t conditional on data for q available at time ($t-T$). Notice that $\hat{q}$ is non-stationary and q_{t-T} follows a random walk.

The second equation is obtained by subtracting expected inflation differentials from UIRP:

$$rid_{t-k} = E_{t-T}\Delta q_{t-k} \tag{A3}$$

where *t*, *T* and *k* represent discrete time and range from zero to infinity; $T = k+1$, i.e. (A3) refers to a one-period maturity bond, and $\Delta q_{t-k} = q_{t-k} - q_{t-k-1}$[9]. Summing up both sides of equation (A3) from $k = 0$ to $k = T$ gives

$$\sum_{k=0}^{T} rid_{t-k} = E_{t-T}\sum_{k=0}^{T}\Delta q_{t-k} \tag{A4}$$

It is possible to show that the sum of *rids* is equal to the temporary component of the real exchange rate by substituting (A4) into (A2) and rearranging, which results

$$\sum_{k=0}^{T} rid_{t-k} = \hat{q} - q_{t-T} \tag{A5}$$

Equation (A5) means that domestic real interest rate rises relative to the foreign when the real exchange rate is temporarily below its equilibrium value and anticipated to grow. From the equilibrium approach to exchange rates as in Stockman (1980) and Stockman

[9] In spite of the fact that (A3) is mathematically derived from UIRP, note that arbitrageurs residents abroad would only care about nominal changes in the exchange rate. On the other hand, transnational agents could care about changes in the real exchange rate as well.

(1988) [for a survey, see Taylor (1995)] is possible to suppose that the incorrect anticipation of a permanent change in the equilibrium real exchange rate arise only if there are variations in **real factors.** .

Substituting (A5) in (A7) and rearranging gives

$$\sum_{k=0}^{T} \varepsilon r_{t-k} + \sum_{k=0}^{T} \varepsilon n_{t-k} = \hat{q} - q_{t-T} \tag{A8}$$

Given long-run money neutrality, nominal factors do not have long-run impacts on the permanent real exchange rate, in other words, when *T* is large or grows to infinity, which implies

$$\sum_{k=0}^{\infty} c_{12}(k) \varepsilon n_{t-k} = 0$$

and concludes the argumentation for the first assumption.

2) rid_t **is not zero**

I assume that rid_t a non null real number i.e., $rid_t \in [-\infty, 0) \vee rid_t \in (0, +\infty]$ which is based on the fact that none of the 107 observations of the sample is equal to zero.

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Figure 2. *Rids*, *Nids and* Inflation Differentials

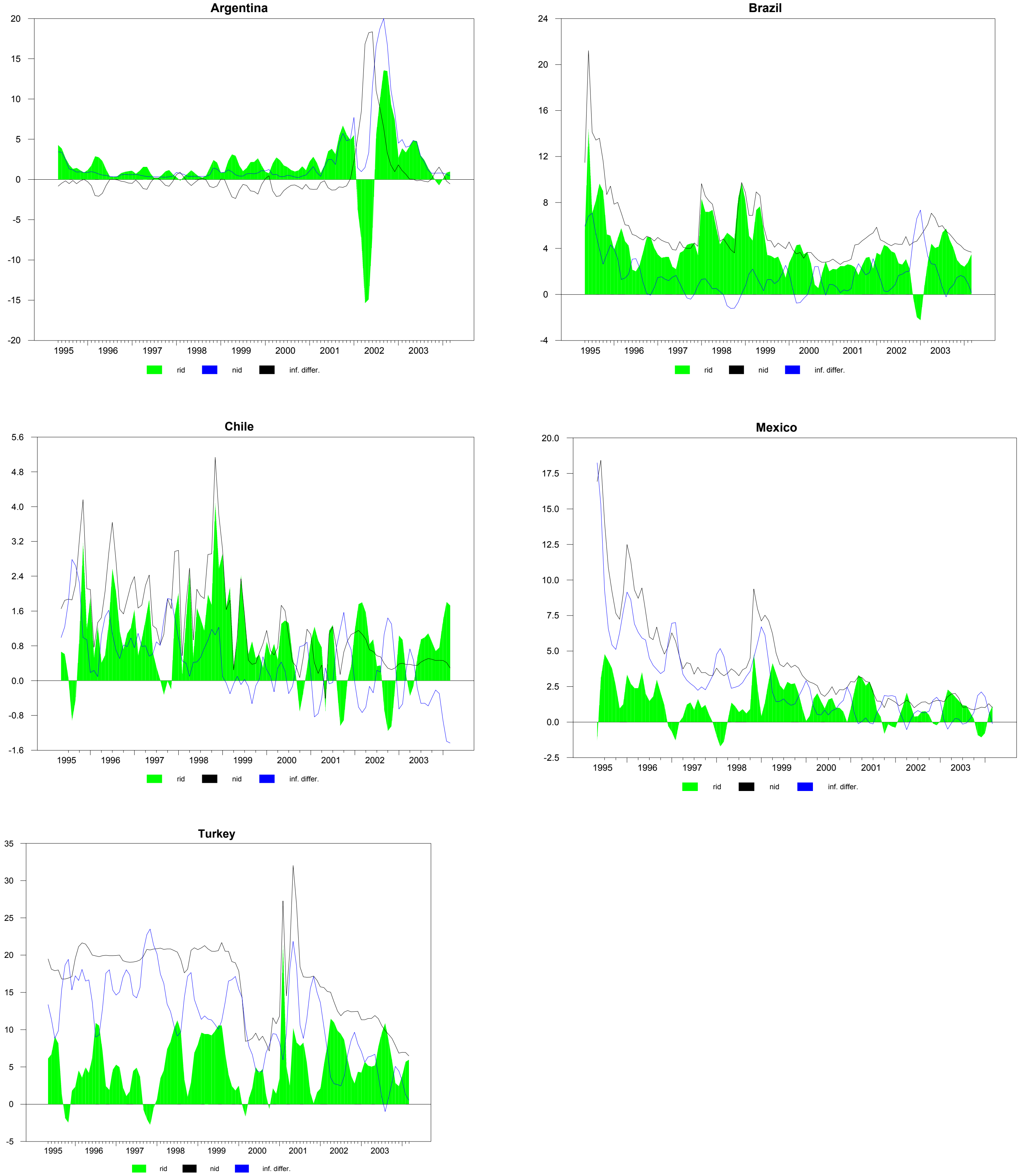
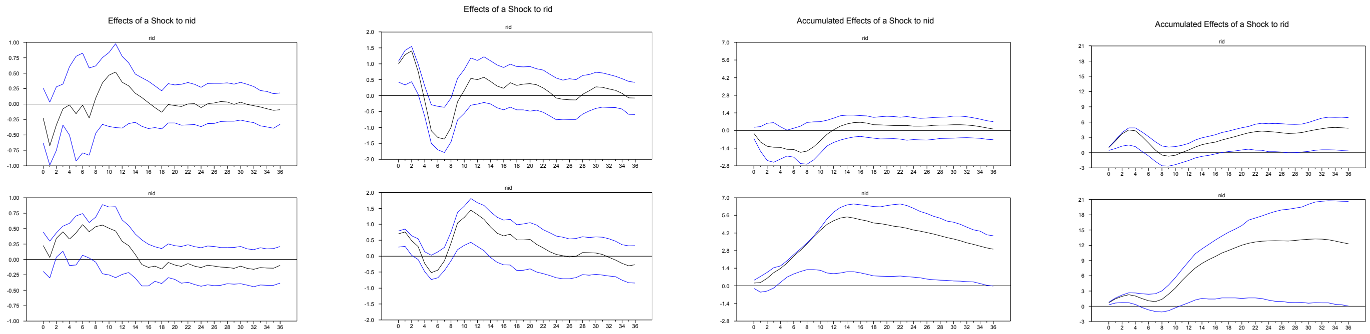
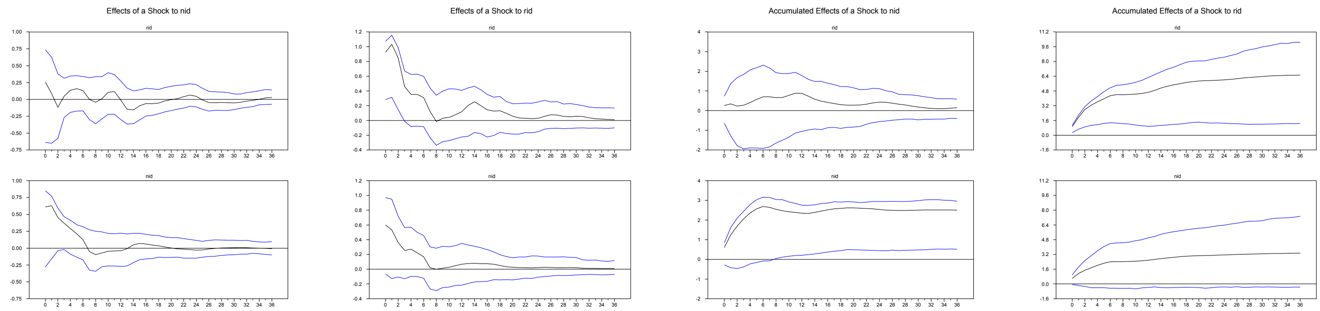


Figure 3. Impulse Responses

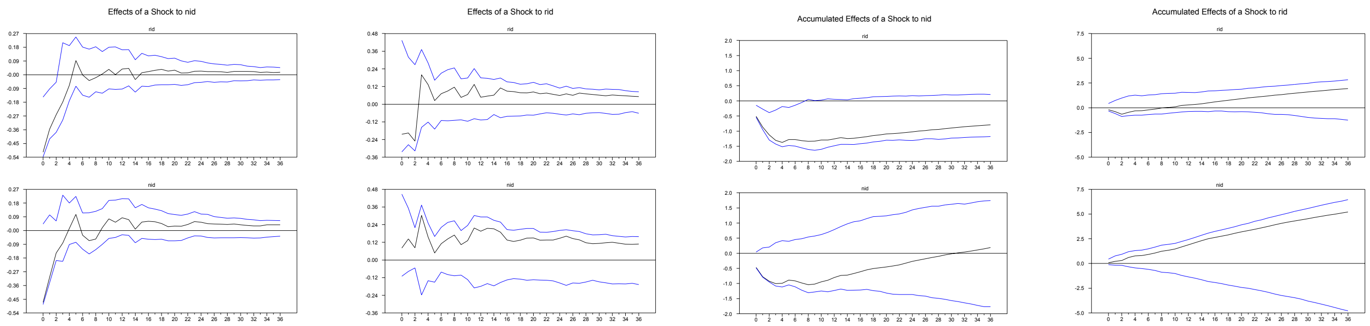
Argentina – US



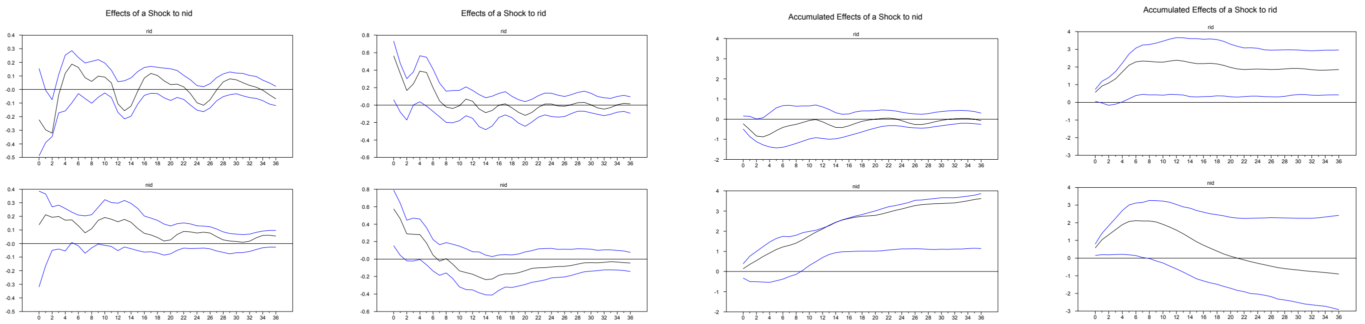
Brazil – US



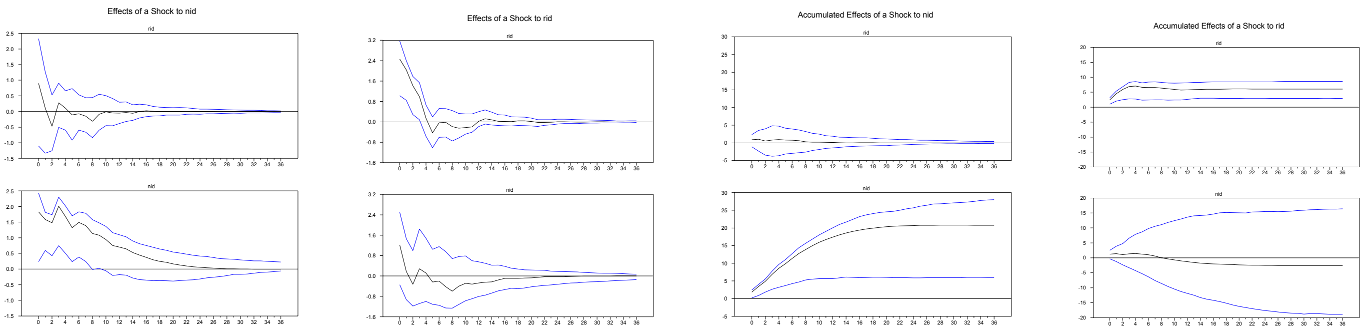
Chile – US



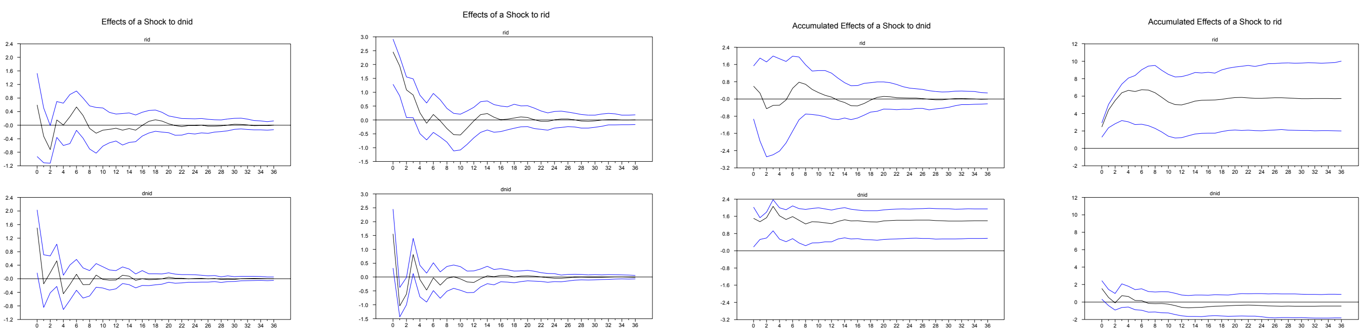
Mexico – US



Turkey – US



Turkey – US (*nids* in first difference)



Notes:
1) Lines in blue represent standard errors which were calculated using one thousand bootstrap draws.
2) The first column shows the impact of a nominal shock while the second column presents the impact of a real shock. The third and fourth columns show the accumulated impact of a nominal and real shock, respectively. *Rids* are on the first line and *nids* on the second.