

Real Interest Parity Decomposition

▪ ALEX LUIZ FERREIRA*

▪ ROSELI DA SILVA**

RESUMO

O objetivo deste artigo é investigar as causas gerais dos diferenciais da taxa de juros real (rids) para um conjunto de países emergentes, para o período de janeiro de 1996 a agosto de 2007. Para tanto, duas metodologias são aplicadas. A primeira consiste em decompor a variância dos rids entre a paridade do poder de compra relativa e a paridade de juros a descoberto e mostra que os diferenciais de inflação são a fonte predominante da variabilidade dos rids; a segunda decompõe os rids e os diferenciais de juros nominais (nids) em choques nominais e reais. Sob certas condições de identificação, modelos autorregressivos bivariados são estimados com tratamento adequado para as quebras estruturais identificadas e as funções de resposta ao impulso e a decomposição da variância dos erros de previsão são obtidas, resultando em evidências favoráveis a que os choques reais são a causa mais provável dos rids.

PALAVRAS-CHAVE

diferencial de juros reais, países emergentes, quebras estruturais, decomposição da variância dos erros de previsão

ABSTRACT

The aim of this paper is to investigate the general causes of real interest rate differentials (rids) for a sample of emerging markets for the period of January 1996 to August 2007. To this end, two methods are applied. The first consists of breaking the variance of rids down into relative purchasing power parity and uncovered interest rate parity and shows that inflation differentials are the main source of rids variation; while the second method breaks down the rids and nominal interest rate differentials (nids) into nominal and real shocks. Bivariate autoregressive models are estimated under particular identification conditions, having been adequately treated for the identified structural breaks. Impulse response functions and error variance decomposition result in real shocks as being the likely cause of rids.

KEYWORDS

real interest rate differentials, emerging markets, structural breaks, breakdown of prediction errors variance

JEL CLASSIFICATION

F32, F36, F21

* Departamento de Economia, FEA-RP/USP. Endereço para contato: Av. Bandeirantes, 3900 – Ribeirão Preto – SP. CEP: 14040-900. E-mail: alexferreira@usp.br.

** Departamento de Economia, FEA-RP/USP. Endereço para contato: Av. Bandeirantes, 3900 – Ribeirão Preto – SP. CEP: 14040-900. E-mail: roselisilva@fearp.usp.br.

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1 INTRODUCTION

Uncovered Interest Rate Parity (UIP) with rational expectations and relative Purchasing Power Parity (PPP) entail the Real Interest Rate Parity Hypothesis (RIPH) [Roll (1979)]. The common finding 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 that *rids* are autoregressive and relatively short-lived (see, for instance, OBSTFELD; TAYLOR, 2003 and GOLDBERG *et al.*, 2003). The aim of the current paper is to investigate the general causes of *rids*. For this purpose, we use a selected sample of emerging markets in which latest evidence indicated that *rids* (in relation to the USA) mean-revert to a positive equilibrium (see FERREIRA; LEÓN-LEDESMA, 2007).

Departures from RIPH can be explained by *ex post* deviations from PPP and UIP. 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 unveiling the causes and understanding their dynamics is essential to design the appropriate macroeconomic policies to change differentials. Finally, the long-run money neutrality is still a motivating question, which is tested in an innovative way.

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 (1998), 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), Camarero and Tamarit (2002) and Jin (2003). Because of Balassa-Samuelson effects, the sign of an impact of a real shock on exchange rates

1 See Iwata and Tanner (2007) for evidence on the trade-off between exchange rate and interest rate volatility in developing countries.

(and *rids*, as we 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 – as observed by MacDonald and Ricci (2004) – *rids* might capture productivity differentials.

We 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 and UIP 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 our 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 draws insights from Levine (1991) and Frankel and MacArthur (1988) but it is based on Cheung *et al.* (2003). The latter has separated the variance of *rids* between deviations from relative PPP and UIP using the relationships given by RIPH as in the following equation

$$rid_t = (i_t - i_t^* - \Delta s_t^e) - (\pi_t - \pi_t^* - \Delta s_t^e) \quad (1)$$

where *rid* is the real interest rate differential, *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 rates of inflation are π_t and π_t^* , respecti-

vely. Observe that $i_t - i_t^* - \Delta s_t^e$ are *ex ante* deviations from UIP and $\pi_t - \pi_t^* - \Delta s_t^e$ correspond to *ex ante* deviations from PPP.

Given the definition of variance and covariance and noting that forecast errors cancel out in (1), we can write

$$\text{Var}(rid_t) = \text{Var}(i_t - i_t^* - \Delta s_t) + \text{Var}(\pi_t - \pi_t^* - \Delta s_t) - 2\text{Cov}(i_t - i_t^* - \Delta s_t, \pi_t - \pi_t^* - \Delta s_t) \quad (2)$$

Another way to decompose the variance of *rids* is by noting that changes in the exchange rate also cancel out in (1). As *rids* are equal to interest rate differentials subtracted from inflation differentials by construction, we can also write

$$\text{Var}(rid_t) = \text{Var}(i_t - i_t^*) + \text{Var}(\pi_t - \pi_t^*) - 2\text{Cov}(i_t - i_t^*, \pi_t - \pi_t^*) \quad (3)$$

As explained by Engel (1996, p. 138), this type of RIPH decomposition “*makes sense – real interest parity could fail either because ex ante PPP fails (goods markets are not integrated) or because uncovered interest parity fails (capital markets are not integrated)*”. Engel (1996) has further criticised the works of Canova (1991), Bekaert (1994), Gokey (1994) and Huang (1990) who decomposed deviations from UIP into deviations from PPP and RIPH because “*Efficiency of the forward market does not require ex ante PPP or ex ante real interest equality. Both could fail, and fail wildly, yet uncovered interest parity could still hold.*” (p. 137). Apart from Cheung *et al.* (2003), the only work performing variance decomposition along the lines set on (2) and (3) is Tanner (1998). However, Tanner’s (1998) paper suffers from the same shortcomings raised by Engel (1996) to the aforementioned previous works. The reason is that Tanner (1998) decomposes both the level and the variance of UIP deviations between deviations from PPP and RIPH.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. In this part of the paper, we base our tests on the methodology employed by Enders and Lee (1997). These authors first ran a VAR using real and nominal exchange rate variations as dependents variables and later applied the Blanchard and Quah (1989) decomposition. They also presented a theoretical model that illustrates the impact of the two types of shocks, real and nominal, on exchange rates. The nominal shock has the property of not affecting the real variable on the long run while there is no restriction for the real shock. From rational expectations UIP,

2 His conclusion for the study for 34 emerging and developed economies is that the variance of *rids* explain most part of the variance of UIP deviations.

we know that there is a theoretical relationship between nominal exchange rate variations and nominal interest rate differentials. This relationship also occurs for real variables, which can be seen by subtracting inflation differentials from UIP as below

$$rid_t = \Delta q_t^e \tag{4}$$

where Δq_t^e represents expected changes in the real exchange rate.. Hence, we can borrow the assumption that real and nominal factors are the disturbances affecting nominal interest rate differentials (*nids*, hereafter) and *rids* from the literature that applied variance decomposition to real exchange rates (see also ROGERS, 1999, and ASTLEY; GARRATT, 2000, for example) and write

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

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

where we ignored intercept terms for simplicity; real and nominal shocks are represented by $\epsilon r_t, \epsilon n_t$ respectively; disturbances are assumed to be iid $N(0, \sigma_\epsilon^2)$ in which σ_ϵ^2 represents variance.

The letter *c* stands for the coefficients associated with the responses of *rids* and *nids* to shocks at each period *k*. The system of equations in (5) and (6) represent an infinite bivariate moving average (BMAR). A BMAR can be represented by a bivariate autoregression model (BVAR) if the roots of the lag polinomials are out of the unit circle, known as the invertibility conditions. The same condition applies to the lag polynomial of the BVAR which guarantees the stability conditions. Under such conditions, the BVAR representation is

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

where e_{1t} and e_{2t} stand for the error terms, which are composite of the pure innovations e_{rt}, e_{nt} .

The Choleski decomposition imposes a contemporaneous restriction in (5) or (6) in order to recover their parameters 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.³ 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.⁴

Another alternative is the method proposed by Blanchard and Quah (1989). For this decomposition we considered that the sum of nominal shocks has a zero impact on the series of *rids*

$$\sum_{k=0}^{\infty} c_{12}(k) \varepsilon n_{t-k} = 0 \quad (8)$$

Following the idea of Faust (1998), as explained below, the restriction in (8) is used to test for robustness of the Choleski decomposition as we cannot think about a theoretical explanation for (8) and recognise its contentious character.

As a matter of fact, 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 BVAR literature, the use of implausible 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.*” (p. 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.*” (p. 210). We 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 We 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.

4 Monetary Policy Committee meetings in Brazil, for example, are realised on a monthly basis and, in most of the cases, interest rates cannot change until the day of the meeting.

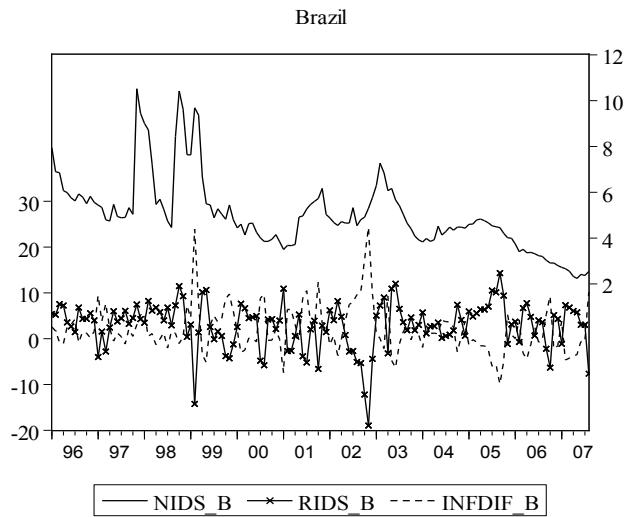
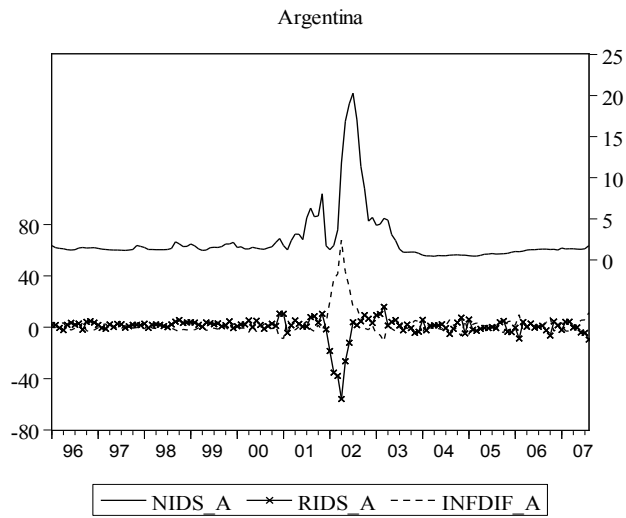
3 RESULTS

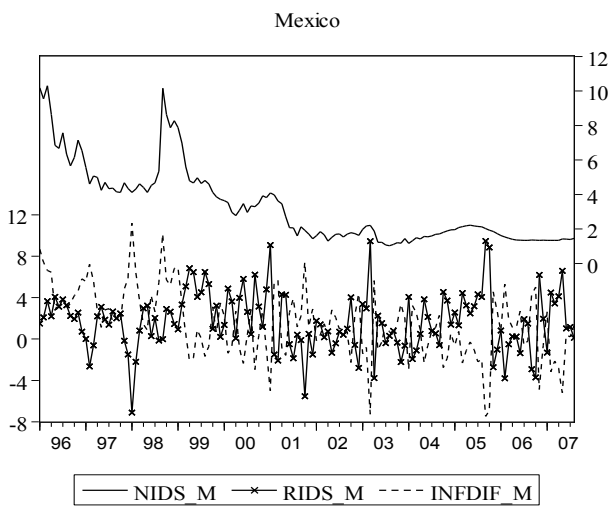
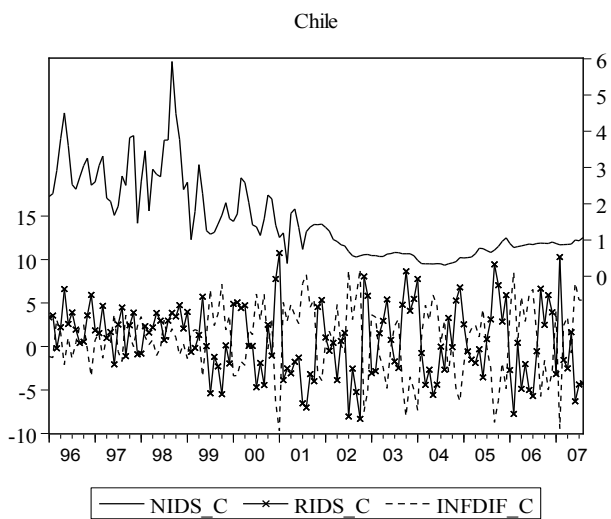
The emerging markets of the sample comprise the small open-economies of Argentina, Brazil, Chile, Mexico and Turkey. We used the USA as the reference large economy for the calculation of the *rid*. The period of the tests corresponds to the interval that spans from 1996M1 to 2007M8.

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. In addition, this period is characterised by various shocks from financial crises: Asian, Russian, Brazilian and Argentinean. The higher volatility that followed these crises justifies the choice for the variance decomposition and innovation accounting. Data on interest rates and average exchange rates was obtained from IMF's International Financial Statistics (IFS). We have chosen the Treasury Bill Rate for Brazil and Mexico while deposit rates for Argentina, Chile and Turkey because of data availability. The inflation rate is either the rate of growth of the Producer Price Index (PPI) or the Wholesale Price Index (WPI), which are more sensitive to variations in the price of tradables. We transformed the annualised monthly interest rate and the inflation rate into compounded quarterly rates and then subtracted the latter from the former. Quarterly exchange rate changes were calculated using the average of the corresponding period.

Graphs are presented in Figure 1 and descriptive statistics of the differentials are reported in Table 1. Note that the means of *rids* are smaller than *nids* for all countries, and the highest *nids* are in Brazil followed by Turkey, Mexico, Argentina and Chile. The higher standard deviations of *rids* is a general result that is driven by the inflation differentials behaviour, as Figure 1 shows. Visual inspection of the time series seems to point out for non-stationarity of Brazil, Mexico, Chile and Turkey *nids*. On the other hand, the outlier of the Argentinean *nids* could lead to the acceptance of a unit root when traditional tests are used; the same would occur for other series that appear to present structural breaks, with the exception of Chile.

FIGURE 1 – RIDS, NIDS AND INFLATION DIFFERENTIALS (INFDIF)





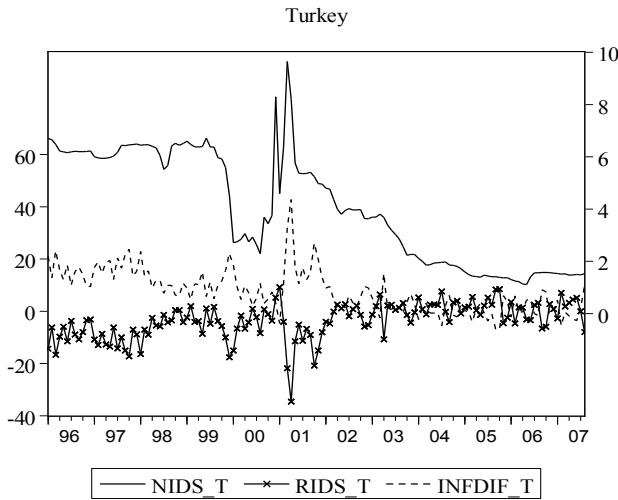


TABLE 1 – SOME DESCRIPTIVE STATISTICS OF RIDS AND NIDS

	Variable	Mean	Min	Max	Std. Dev.
Argentina	<i>nids</i>	2.34	0.44	20.28	3.29
	<i>rids</i>	0.31	-55.96	15.93	8.18
Brazil	<i>nids</i>	5.01	2.24	10.52	1.62
	<i>rids</i>	2.98	-18.98	14.30	5.05
Chile	<i>nids</i>	1.53	0.30	5.92	1.09
	<i>rids</i>	0.74	-8.31	10.75	4.07
Mexico	<i>nids</i>	3.38	1.04	10.30	2.22
	<i>rids</i>	1.70	-7.10	9.48	2.83
Turkey	<i>nids</i>	4.04	1.11	9.64	2.12
	<i>rids</i>	-3.55	-34.56	9.21	6.87

In order to find out the order of integration of *nids* before running the BVAR, we initially tested for the presence of unit roots.⁵ Considering the low power problems and size distortions of the traditional tests (Augmented Dickey-Fuller - ADF, Phillips-Perron - PP - and Kwiatkowski, Phillips, Schmidt and Shin - KPSS), largely pointed in the literature⁶, we applied more robust tests such as the DF-GLS (ELLIOT *et al*, 1996 and ELLIOTT, 1999) and Ng-Perron (1996, 2001) tests.

⁵ We have not performed unit root tests for *rids* as this has already been done for the countries of our sample, see for example Ferreira and León-Ledesma (2007). As pointed out earlier, results show that this variable is stationary.

⁶ See Maddala and Kim (2003) for a survey of the literature.

We first apply DF-GLS (ELLIOTT; ROTHENBERG; STOCK, 1996) who propose a simple modification to the ADF, in which the time series is previously filtered from its deterministic components. The first set of tests, which goes along the lines of the ADF, allows for an adequate analysis of the series in the presence of deterministic components.

The second test, proposed by Ng and Perron (1996, 2001), follows the non-parametric methodology of the PP tests, in which the variance-covariance matrix of the parameters from the test equation is robust to heteroscedasticity and autocorrelation. The authors propose to treat the problems associated with the usual tests by building a test statistic without the deterministic components (the latter estimated by GLS) and spectral density function at zero frequency estimated as an AR(1) process (PERRON; NG, 1998).

We found the optimal augmentation lag using a Modified Akaike information Criterion (MAIC), following Ng-Perron (2001). We report the deterministic components, lag specification, t-ratios, and critical values for DF-GLS and Ng-Perron tests in Table 2.

TABLE 2 – UNIT ROOT TESTS ON NIDS

	Specification		Test Statistic	
	Lags ¹	Determinist	DF-GLS	Ng-Perron
Argentina	5	Constant	-1.98*	-7.71**
Brazil	0	Constant, Linear trend	-3.69*	-22.71*
Chile	17	Constant, Linear trend	-0.99	-1.39
Mexico	0	Constant, Linear trend	-1.77	-6.33
Turkey	2	Constant, Linear trend	-2.48	-11.81

Notes: 1 Starting from 13 lags (except Chile, 24), MAIC Selection.

* Rejection of the null at 5% (** at 10%) confidence level.

With the exception of Argentina, the most adequate test specification, according to the behavior of the residuals, contains a deterministic trend. The most complicated deterministic behavior to be modeled is Chile, which seems to require a non linear specification. This is shown by the excessive number of lags selected by the MAIC. Table 2 shows that both tests rejected the null of a unit root for Argentina and Brazil, but did not reject the null of the unit root for Chile, Mexico and Turkey. Graphical analysis in Figure 1 suggests the existence of breaks in the series, as pointed before. It is known that such breaks, as well as the presence of outliers, can generate a bias on the standard unit root tests leading to a false rejection of null hypothesis (MADDALA; KIM, 2003). In order to investigate for this possibility,

we performed unit root tests that account for two possible structural breaks: the Lee and Strazicich (2003) test, LS test hereafter.

The advantage of the LS test, besides the endogenous investigation of two possible breaks, is to specify a test with breaks on both the null and the alternative which does not leave any ambiguity regarding the trend in the series: the rejection of the null implies a trend-stationary series. Suppose the following data generation process:

$$nid_t = \delta'Z_t + e_t \quad (9)$$

$$e_t = \beta e_{t-1} + \varepsilon_t \quad (10)$$

Where Z_t is a vector of exogenous variables and $\varepsilon_t \sim iid N(\mathbf{0}, \sigma^2)$, δ is a vector of parameters. Model A (crash model) allows for two breaks in level, including two dummies D_{1t} and D_{2t} , so $Z_t = [1, t, D_{1t}, D_{2t}]$, where $D_{jt} = 1$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and 0 otherwise, and model C includes two breaks in level and trend, a changing growth model where $Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]$ where $DT_{jt} = 1$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and 0 otherwise. The results are reported in Table 3.

The date breaks retrieved by the tests seem to reflect the effects of domestic crises in Argentina and Brazil. Turkey is the only case that had just one significant break, which coincides with the falling trend of its nominal interest rate differentials. The exogenous events most closely associated with the date break of Mexico and Chile are the Asian and the Brazilian financial crisis, respectively. Chilean *nids* presented relative stability from 2001 onwards, inverting its falling trend in the middle of 2004, which corresponds to another date break. Inspection of Table 3 shows that the unit root was strongly rejected for Argentina, Chile, Mexico and Turkey. In the last three cases, the results reveal that the former unit root tests did not mistakenly rejected the null hypothesis even in the presence of breaks in the deterministic trends. On the other hand, tests for Brazil reveal some ambiguity: without considering the breaks the series did not present a unit root; by considering them, one could not reject the unit root. This result is taken into account later on the BVAR specification.

TABLE 3 – NIDS: TWO-BREAK MINIMUM LM UNIT ROOT TEST (LS TEST)

	Optimal lag	10% Significant level break dates		Test statistic	Critical value break points
Argentina	16	2002M1 (level; trend)	2003M9 (level; trend)	-8.1326 ***	$\lambda = (0.52, 0.66)$
Brazil	1	1999M3 (level)	2001M6 (trend)	-5.1777	$\lambda = (0.28, 0.47)$
Chile	14	1998M11 (level)	2004M6 (trend)	-6.7446 ***	$\lambda = (0.25, 0.73)$
Mexico	18	1998M7 (level; trend)	2001M11 (trend)	-9.7997 ***	$\lambda = (0.22, 0.51)$
Turkey	10	2001M5 (level; trend)	2004M3 (none)	-6.2670 ***	$\lambda = (0.46)$

Obs: critical values are shown below for the one and two-break minimum LM unit root test with linear trend (Model C) at the 1%, 5%, and 10% levels for a sample of size $T \frac{1}{4} 100$, respectively.

Critical values are symmetric around λ and $(1 - \lambda)$, where $\lambda_i = T\lambda_i/T$ and can be interpolate at additional break points.

Notes: *, **, *** significant at the 10%, 5%, and 1% levels, respectively.

Obs: the critical values shown below come from Table 2 in Lee and Strazicich (2003) for two breaks and from Strazicich *et al.* (2004) for one break.

	Two-breaks critical values			One-break critical values			
	1%	5%	10%	1%	5%	10%	
$\lambda = (0.2, 0.4)$	-6.16	-5.59	-5.27	$\lambda = (0.4)$	-5.05	-4.50	-4.18
$\lambda = (0.2, 0.6)$	-6.41	-5.74	-5.32	$\lambda = (0.5)$	-5.11	-4.51	-4.17
$\lambda = (0.2, 0.8)$	-6.33	-5.71	-5.33				
$\lambda = (0.4, 0.6)$	-6.45	-5.67	-5.31				
$\lambda = (0.4, 0.8)$	-6.42	-5.65	-5.32				
$\lambda = (0.6, 0.8)$	-6.32	-5.73	-5.32				

Regarding variance decomposition, results presented in Table 4 demonstrate that the share of *ex post* deviations from UIP in the variance of *rids* is higher than the share of *ex post* deviations from PPP for Argentina, Brazil, Chile and Turkey.

TABLE 4 – VARIANCE DECOMPOSITION OF RIDS BETWEEN UIP AND PPP DEVIATIONS

	Argentina	Brazil	Chile	Mexico	Turkey
Variance of:					
<i>Rids</i>	66.9	25.5	16.6	8.0	47.2
Deviations from UIP	306.5	98.4	19.6	14.5	122.3
Deviations from PPP	122.4	68.7	14.4	17.2	72.6
% of <i>Rids</i>' variance:					
Deviations from UIP	458.3	386.4	118.1	180.8	259.3
Deviations from PPP	183.0	269.8	87.2	215.3	153.9
-2cov(UIP,PPP)	-537.5	-552.2	-104.6	-293.9	-311.0

The high volatility of exchange rates is responsible for most part of the variance of the 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 inflation differentials are the predominant source of variability for most *rids* of the sample. The variance of *nids* is higher during the period of the financial crises during the 1990s. After the nineties they are much more stable and, for this reason, the variance of the inflation differential dominates the series.

TABLE 5 – VARIANCE DECOMPOSITION OF RIDS BETWEEN NIDS AND INFLATION DIFFERENTIALS

	Argentina	Brazil	Chile	Mexico	Turkey
Variance of:					
<i>Rids</i>	66.9	25.5	16.6	8.0	47.2
<i>Nids</i>	10.8	2.6	1.2	4.9	4.5
Inflation differential	88.7	26.3	15.4	10.8	68.5
% of <i>Rids</i>' variance					
<i>Nids</i>	16.2	10.4	7.2	61.5	9.5
Inflation differential	132.6	103.1	92.9	135.5	145.3
-2cov(<i>Nids</i> , Inf. Differential)	-48.5	-13.4	-0.1	-96.4	-54.4

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. In conclusion, the volatility of inflation differentials explains *rids*' variance.

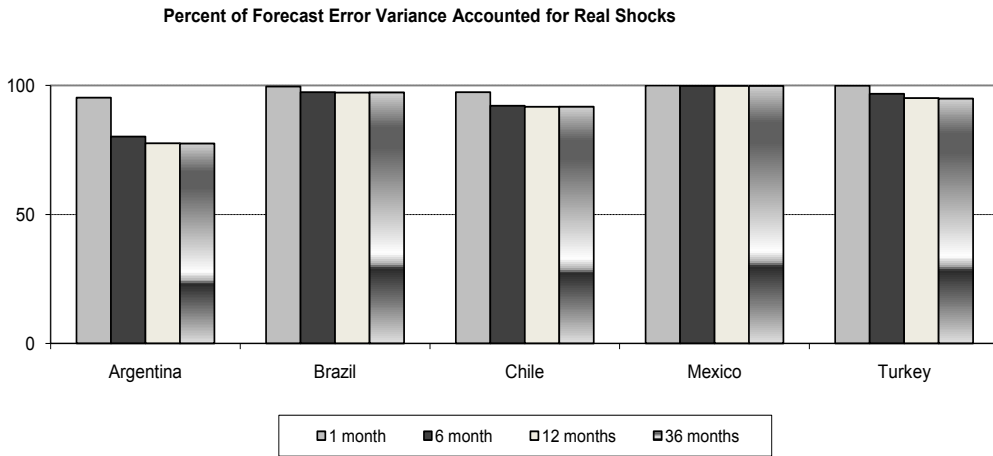
We turn to the findings of innovation accounting by first analysing forecast error variance decompositions. For this purpose, we estimated a BVAR for each country according to equation (7). However, the results of trend-stationary *nids* indicated that there are deterministic trend components, although with breaks (Tables 2 and 3), which must be taken into account in order to guarantee a stable BVAR. As we are interested in decomposing the error variance of the BVAR, we need to specify it in a way that errors are white-noise with the stationarity conditions met. Table 6 summarise the results.

TABLE 6 – BVAR SPECIFICATION

	Deterministic Components (following UR tests evidences)	Lags	Greatest Root of Characteristic Polynomial (Modulus)
Argentina	Constant, trend and break dummies level (1 and 2) and trend (1 and 2)	5	0.8876
Brazil	Constant, trend and break dummies level (2) and trend (1 and 2)	1	0.6382
Chile	Constant, trend and break dummies level (2) and trend (1 and 2)	4	0.7569
Mexico	Constant, trend and break dummies level (1 and 2) and trend (1 and 2)	1	0.6778
Turkey	Constant, trend and break dummies level (1 and 2) and trend (1 and 2)	5	0.8269

Figure 2 shows the percentage of variance explained by real shocks for some selected time-horizons: 1, 6, 12 and 36 for Choleski decomposition. Real shocks are the main source of variation in *rids* for all countries at all horizons according to the Choleski decomposition.

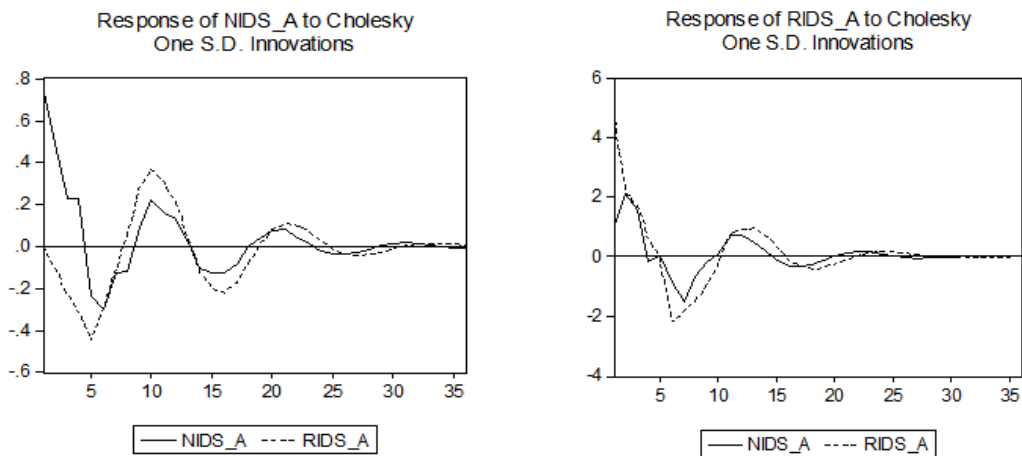
FIGURE 2 – FORECAST ERROR VARIANCE DECOMPOSITION OF RIDS



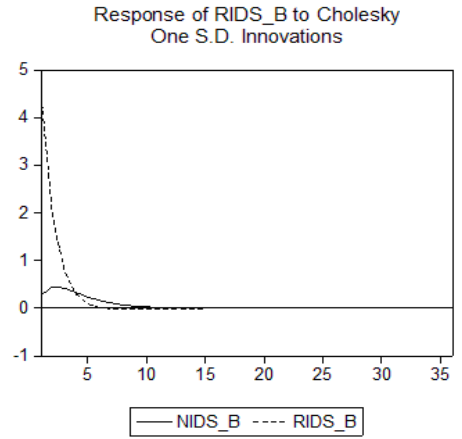
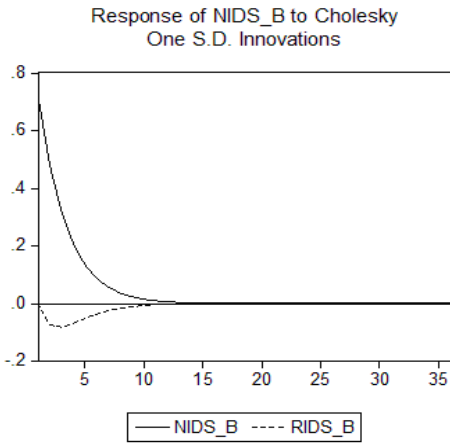
Note: The forecast error variance decomposition is the percentage of the mean squared error due to a real shock.

Figure 3 presents impulse responses obtained through the use of Choleski technique as short-run responses would be somewhat influenced by the lagged restriction. Long run restrictions leave the short run dynamics of the BVAR 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.

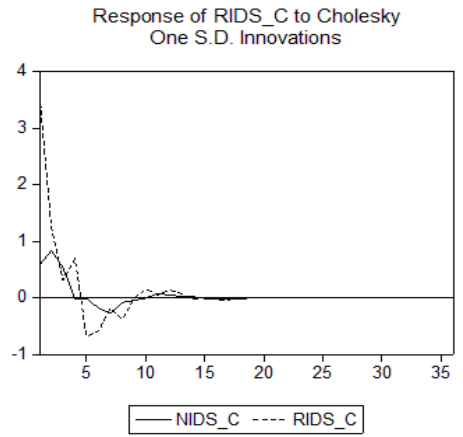
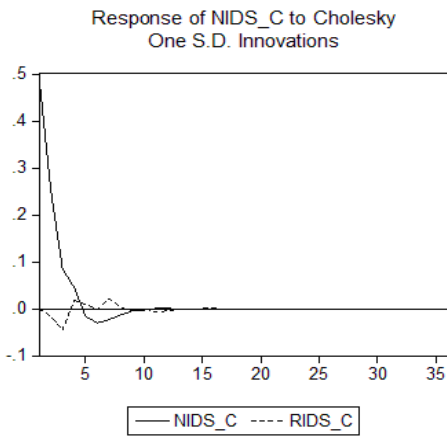
FIGURE 3 – IMPULSE RESPONSES – ARGENTINA



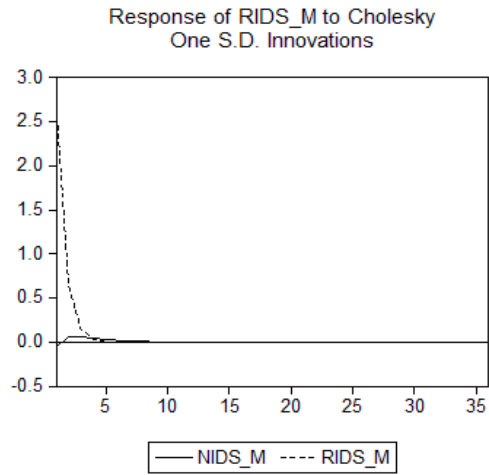
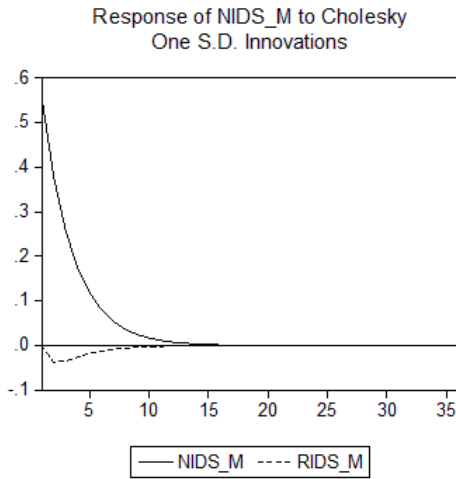
BRAZIL



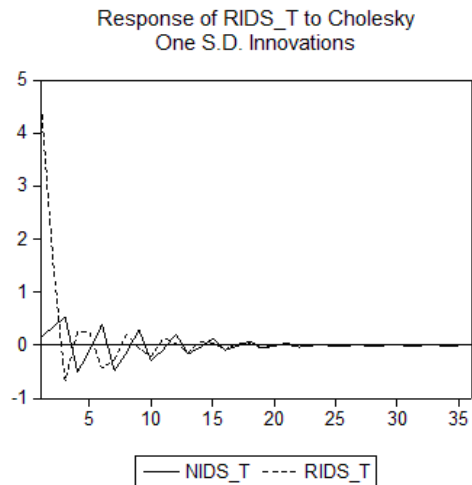
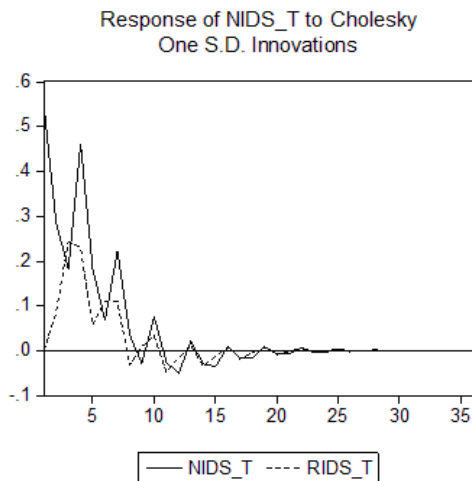
CHILE



MEXICO



TURKEY



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 raises *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, both *rids* and *nids* of Argentina react in a similar way to either a real or a nominal shock. The oscillation pattern is the same for both *rids* and *nids*. A nominal shock has an initial positive impact (until the 5th period) over *nids* while the real shock has a negative impact. The response of Argentinean *rids* and *nids* to nominal and real shocks follow the same pattern, with positive effects until the 5th period and a change of signs every five periods (approximately).

The behaviour of the impulse response function for Turkey is similar to the one of Argentina. It oscillates and the convergence is slow, which is a result of a longer short run dynamics (both Turkey and Argentina were found to have 5 optimal augmentation lags, see Table 6). However, this is the only country in which *nids* respond positively to a real shock, which occurs from the second lag onwards.

Impulse response for Brazil and Mexico converge exponentially. This result is in accordance to its short-run dynamics, which presents one lag. *Nids* respond to a real shock negatively and converge to zero from the third period onwards. Chile has a more complex short-run dynamics, (the optimal augmentation lag is 4) and presents an undershooting of both *nids* and *rids* in response to a nominal shock.

As show in Table 7, a real shock has a negative effect on *nids* for all countries, with the exception of Turkey. For those countries a nominal shock leads to a strong long run impact over inflation rather than to *nids* (for Chile this effect is small). A one standard deviation shock increases *nids* of all countries more than proportionally, with Chile being the exception. The results for Chile could be associated to the fact that, among all countries in our analysis, this is the economy that has the oldest and most successful stabilisation process. Structural reforms in financial markets and in the fiscal regime were done in the nineties with gradual adoption of an inflation target regime. On the other hand, *rids* of all countries present positive accumulated effects and, as expected by the identifying restriction, they are greater for a real shock rather than to a nominal one. The reason for a higher accumulated impact than the initial increase might be related to frictions in goods, financial markets or to the breakdown of rational expectations that cause the series to present dynamics.

TABLE 7 – ACCUMULATED IMPULSE RESPONSE FUNCTIONS

	<i>Nids</i>			<i>Rids</i>		
	Maximum Convergence lag	Nominal	Real	Maximum Convergence lag	Nominal	Real
Argentina		1.23	-0.75		2.80	5.10
Brazil	17	2.09	-0.38	16	2.22	7.50
Chile	20	0.79	-0.02	22	1.45	4.33
Mexico	20	1.74	-0.15	11	0.23	3.49
Turkey	25	1.91	0.75	31	0.27	5.57

Finally, while the sign of the accumulated impact of real shocks on *nids* is negative on average, they are positive for *rids* of all countries. As the 1990's was a period characterised for productivity increases, this result, *prima facie*, lends support for Balassa-Samuelson effects.⁷ However, the 1990's was also plagued by financial crisis which possibly imply risk premium shocks.

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 inflation differentials explains most part of the volatility of *rids* for all countries. 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.

We found evidence of trend-stationarity for most *nids* in our sample. 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 accumulated impact of a real shocks on *nids* is negative while it is positive for *rids* of all countries. At the extent to which real sho-

⁷ See Lee and Tang (2007) for latest survey and evidence on the relationship between productivity and real exchange rates. See Faria and Leon-Ledesma (2003) for a test of Balassa-Samuelson effects on developed countries.

cks reflect productivity changes, this result provides support for Balassa-Samuelson effects. However, the 1990s was a period of various financial crises and the results of endogenous date breaks seem to reflect this fact. Finally, nominal shocks impact positively on *rids* and *nids* in the long-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 real shocks in the path of *rids* and points out to deviations from UIP as their driving source. *Nids* were found to be trend stationary, probably reflecting the tendency in reduction following the financial crises in the 1900s. The conclusion is that one should look at unexpected productivity changes and risk premium shocks in order to comprehend the dynamic behaviour of real differences in returns across countries.

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