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The Simultaneity Bias of the Uncovered Interest Rate Parity: evidence using survey data for Brazil

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Abstract

The paper focuses on testing the ex ante uncovered interest rate parity (UIP) in a single equation framework. I use survey based expectations as a proxy for observed exchange rate expectations and IV estimation for statistical inference, overcoming the negative bias indicated by the small scale general equilibrium model with standard ingredients. Results from 2001M11 until 2014M12 using data for Brazil and the United States of America (Real/Dollar exchange rate) show strong support for ex ante UIP.

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1 Introduction

Although support for uncovered interest rate parity (UIP, hereafter) has been growing, this hypothesis still causes embarrassment from the empirical point of view[1]. There are competing explanations for the failure of short-run UIP (for example, risk, Peso problems, improper econometric techniques etc) but none seem to be widely accepted and there is no consensus on the subject. Hence, there is an open field for investigation and space to work towards some sort of consensual explanation.

This paper builds on the work of McCallum (1994b) who puts forward a model that recognizes the simultaneous action of agents and Central Bankers in determining equilibrium interest and exchange rates. His model assumes interest rate smoothing and reaction against exchange rate changes, and implies that the failure of the hypothesis can be policy driven.

The endogeneity issue has already been investigated. Many authors have recognized the potential of this explanation in solving the UIP problem, for example, Meredith and Chinn (1998) and Favero and Giavazzi (2004) (to cite just a few) and for exchange rates, in particular, one can see Engel and West (2005). Kugler (2000) was a pioneer in noting the main implications of McCallum (1994b)'s model for Ordinary Least Squares (OLS) estimations. He applied the model to analyze the term structure of interest rates and derived the asymptotic bias using McCallum (1994b)'s policy reaction function. On the other hand, Christensen (2000) tested the policy reaction function of the McCallum (1994b) model for the US, Germany and Japan but did not find supporting evidence regarding the size of the parameters needed to generate the negative bias.

Our contribution is twofold. First, we test UIP taking into account the endogeneity problem. Second, our tests are performed using market data on exchange rate expectations collected from the unique - as it is publicly available - database of the Brazilian Central Bank from 2001M11 to 2014M12. Brazil is also an interesting case, because it has been experiencing high nominal interest rates compared to other developed economies, at least during our sample period.

On the theoretical side, we develop a simple macroeconomic model that does not hinge on the assumption of "leaning against the wind" as McCallum (1994b). However, it shows that reaction against prices can be enough to generate the bias on UIP. Furthermore, we show the associated asymptotic bias and provide a hint as to the size of the structural parameters needed to generate a negative bias. Our results show that Instrumental Variable (IV) estimation of *ex ante* UIP reduces the OLS bias. There is evidence supporting *ex ante* UIP and that the dynamically complete model, which better represents the observed data (when variables are in equilibrium), produces the best fit.

The rest of the paper is organized as follows. We first present the model and discuss both the asymptotic bias and the fully dynamic model for expected exchange rate changes. The penultimate section is dedicated to the empirical findings and the final to conclusions.

2 Endogeneity

We deduce from McCallum (1994b)'s article that the empirical failure of short-run UIP is due to researchers overlooking the fact that this hypothesis, concerning equilibrium in the assets market, belongs to a system of equations. Hence, a regression of UIP using OLS produces estimated parameters that cannot have a structural interpretation and could also be subject to simultaneity bias. A shortcoming of his model is that it contained only two equations, the policy function and UIP itself. Another complication is that monetary authorities react to exchange rate changes but not to deviations of inflation from its target[2]. In order to overcome these limitations, our model considers a Taylor rule type function under a strict inflation target as well as other equilibrium relationships, such as the modeling of the demand and the supply side of the economy.

This section aims to illustrate how a negative bias can arise from the OLS regression. Our objective is to obtain a closed-form analytical solution for the reduced form model along the lines of McCallum (1994b) and Engel and West (2005), for instance, but without resorting to the explicit inclusion of exchange rates in the policy function. A possible justification is the non supportive result presented by Christensen (2000) on

[1] See, for example, Isard (2006) and Chinn (2006).

[2] As a matter of fact, McCallum (1994a) had already recognized that a more general reaction function - including reaction against prices - would be theoretically plausible and could be empirically stronger.

"leaning against the wind". The model presented here describes a simplified open economy, as opposed to others using more detailed model specifications - see for instance, the interesting works of Meredith and Ma (2002) and Alexius (2002). A complex structure would require numerical solutions and simulations, which is an avenue of investigation that we chose not to follow.

As can be seen below, the first equation of the system stands for the UIP relationship under imperfect capital mobility, while the remaining three equations represent the monetary policy reaction function, the Phillips curve and the IS relationship, respectively. As can also be inferred, they result from the subtraction of the foreign equation from the domestic counterpart, assuming that parameters are analogous in both economies:

$$s_t = s^e_{t+1} - (i_t - i^*_t) + \xi_t, \tag{1}$$

$$i_t - i^*_t = \rho(i_{t-1} - i^*_{t-1}) + (1-\rho)[i^n_t - i^{*n}_t + \lambda[\pi_t - \pi^*_t - (\pi^T - \pi^{*T})], \tag{2}$$

$$\pi_t - \pi^*_t = \phi\Delta s_t + (1-\phi)(\pi^T - \pi^{*T}) + \eta_1(h_t - h^*_t) + e^s_t, \tag{3}$$

$$h_t - h^*_t = -\eta_2[i_{t-1} - i^*_{t-1} - (\pi_t - \pi^*_t)] + e^d_t, \tag{4}$$

where s_t is the natural logarithm of the nominal exchange rate, defined as the domestic price of the foreign currency; i_t is the nominal interest rate paid on a one-period bond. The superscript e denotes expected values and the asterisk denotes an exogenous determined foreign variable or the foreign economy; ξ_t represents all other variables that explain differences in nominal returns. One can think of ξ_t as a risk term, which is often done in the literature. For simplification, we start with the assumption that ξ_t is white-noise. The variable π_t stands for the inflation between $t-1$ and t and π^T is the inflation target for $t+1$ known at t; π^T will be constant and equal to zero by hypothesis, i.e. $\pi^T = \pi^{*T} = 0$. The variable i^n_t is the neutral interest rate, i.e. $i^n_t = r_t + \pi^T$. The letter r_t represents the long-run real interest rate, which is determined by real factors such as the marginal product of capital adjusted for risk. The time subscript in r is explained by the time-varying marginal product of capital, which implies a time-varying, neutral, real rate. The log of the output gap is represented by h_t. Error terms e^s_t and e^d_t stand for supply and demand shocks, respectively, and are both random variables. The other letters are parameters: ρ is the smoothing term, $0 < \rho < 1$; λ measures the extent to which money authorities react to deviations of inflation from target, and $\lambda > 1$; η_1 and η_2, both positive quantities, measure the sensitivity of the actual inflation differential to the output gap and the sensitivity of the output gap to the lagged real interest rate, respectively. Our *ad hoc* Phillips Curve equation, (3), was assumed to simplify matters. The parameter ϕ can be thought of as showing the extent to which the inflation differential is anchored in relative purchasing power parity and $(1-\phi)$ on the inflation target differential, and $0 < \phi < 1$. One could also think of ϕ as measuring some sort of pass-through mechanism and $1-\phi$ the degree of credibility of the Central Bank. Also note that we can write

$$i^n_t - i^{*n}_t = r_t + \pi^T - (r^*_t + \pi^{*T}).$$

Hence, the process for the nominal natural interest rate differential is simply given by the real interest rate differential which we express, by assumption, as

$$i^n_t - i^{*n}_t = r_t - r^*_t = \xi_t + \mu_t, \tag{5}$$

where μ_t is an error term; the previous hypothesis can be reasonable if the long-run real interest rate differential is given by the differences in the marginal product of capital between the small and the larger risk-free economy, which is assumed to be equal to the risk premium. The meaning of (5) is that, in the absence of shocks and in initial equilibrium, the monetary authority will set the nominal interest rate at a level equal to ξ_t when $\rho = 0$, that will not induce flows of capital. In other words, the rule prescribes adjusting i_t to shocks in risk at a fraction $1-\rho$.

As UIP is often tested using $nid_t = \Delta s^e_{t+1} + \xi_t$ with some proxy for Δs^e_{t+1}, where Δ stands for the first difference and $nid_t = i_t - i^*_t$, we show that nid_t and ξ_t are correlated. We start by substituting (4) into (3) which gives

$$\pi_t - \pi_t^* = \phi\Delta s_t + \eta_1\{-\eta_2[i_{t-1} - i_{t-1}^* - (\pi_t - \pi_t^*)] + e_t^d\} + e_t^s. \tag{6}$$

Substituting (5) and (6) into (2) and solving the resulting expression for the nid_t, we can write

$$nid_t = \alpha_0 nid_{t-1} + \alpha_1\Delta s_t + \alpha_2\xi_t + e_t, \tag{7}$$

where

$$\alpha_0 \equiv \frac{[\lambda(1-\rho)+\rho]\eta_1\eta_2 - \rho}{\eta_1\eta_2 - 1},$$

$$\alpha_1 \equiv \frac{\lambda\phi(\rho-1)}{\eta_1\eta_2 - 1},$$

$$\alpha_2 \equiv 1 - \rho,$$

and,

$$e_t \equiv \frac{(\rho-1)[\lambda(e_t^s + \eta_1 e^d) + (1-\eta_1\eta_2)\mu_t]}{\eta_1\eta_2 - 1}.$$

Observe that the variable nid_{t-1} is predetermined and, because ξ_t, e_t^s, e^d and μ_t are all i.i.d., e_t is also exogenous and i.i.d. In order to obtain the reduced form, we take into consideration rational expectations UIP by substituting the process for the nid_t in (7) into equation (1) and solving for expected exchange rate changes

$$\Delta s_{t+1}^e = \alpha_0 nid_{t-1} + \alpha_1\Delta s_t + (\alpha_2 - 1)\xi_t + e_t. \tag{8}$$

Then we postulate a bubble-free linear solution using the relevant state variables, as below

$$\Delta s_t = \gamma_0 nid_{t-1} + \gamma_1\xi_t + \gamma_2 e_t. \tag{9}$$

In order to solve for Δs_t, we use the method of undetermined coefficients. After abandoning a non-stationary root ($\gamma_0 = 1$), one reaches the following solution for the nid_t.

$$nid_t = \frac{\lambda\phi(\rho-1)}{\rho(\lambda\phi-1) + [\lambda + \rho(1-\lambda)]\eta_1\eta_2 - \lambda\phi}\xi_t \tag{10}$$

The numerator in (10) is likely to be different from zero. Although there is disagreement in the literature regarding the size of the smoothing parameter, it is probably larger than zero and different from one, for our smaller frequency data (see, for instance, the discussion in Coibion and Gorodnichenko (2012) and Rudebusch (2006)). Also, the λ coefficient, which represents reaction against price changes, is certainly greater than 0 and probably larger than one[3]. Finally, there is an extensive amount of literature documenting that the degree of exchange rate pass-through, although certainly not large, especially for developed economies, is greater than zero[4]. In summary, both the available evidence and our theoretical model allow us to conclude that the nominal interest rate differential and the variable ξ_t are very likely to be correlated. In that case, OLS will render biased and inconsistent estimators.

2.1 Asymptotic Bias

Now, if you wish to estimate equation (1) by OLS

$$\Delta s_{t+1}^e = \beta_0 + \beta_1 nid_t + \epsilon_t \tag{11}$$

where $\epsilon_t \equiv -\xi_t$. The asymptotic value of β_1 will be

[3]Taylor (1995), for instance, reports a value close to 1.5, for the case of $\rho = 0$, as the interest rule cannot prescribe a passive behavior to above the target changes in inflation.

[4]NogueiraJr. and León-Ledesma (2009), for instance, estimated a significant long-run pass-through coefficient of approximately 8% for Brazil during the Inflation Targeting Regime which started in 1999M1 (their sample spans from 1995M1 to 2007M12).

$$\text{plim}(\hat{\beta}_1) = \beta_1 + \frac{\text{Cov}(nid_t, \epsilon_t)}{\text{Var}(nid_t)}. \tag{12}$$

where plim is the probability limit when t grows to infinity. Hence, Bias $= \text{Cov}(nid_t, \epsilon_t)/\text{Var}(nid_t)$. As $\text{Var}(nid_t) > 0$, the sign of the bias will depend on how $\text{Cov}(nid_t, \xi_t)$ differs from zero. As $\epsilon_t = -\xi_t$, the bias will be negative only if $\text{Cov}(nid_t, \xi_t) > 0$[5]. Given the assumption of a zero mean, serially uncorrelated ξ_t, we can write

$$\text{Cov}(nid_t, -\xi_t) = -\mathbb{E}(nid_t \xi_t), \tag{13}$$

and, hence, we need to find $\mathbb{E}(nid_t \xi_t)$, as shown below

$$\mathbb{E}(nid_t \xi_t) = \mathbb{E}\left\{ \frac{\lambda\phi(\rho - 1)}{\rho(\lambda\phi - 1) + [\lambda + \rho(1-\lambda)]\eta_1\eta_2 - \lambda\phi} \xi_t^2 \right\}.$$

Since $\lambda\phi > 0$ and $0 < \rho < 1$, the numerator of the expression above will be negative. It can be shown that, for reasonable parameter values, that the denominator will generate a negative covariance, leading to statistical bias.

Since the β_1 of the "population" is equal to one, the estimated parameter can be negatively biased according to the simple model above. This suggests, for instance, that IV estimation is more appropriate for UIP tests, provided one has proper instruments. However, as the structural equation will not represent the observed data for the expected exchange rate change, we will first derive its reduced form, dynamically complete model.

The dynamically complete model

Substituting the process for nid_t from (10) into (1) and solving for expected exchange rate changes gives

$$\Delta s_{t+1}^e = (\kappa - 1)\xi_t, \tag{14}$$

where

$$\kappa = \frac{\lambda\phi(\rho - 1)}{\rho(\lambda\phi - 1) + [\lambda + \rho(1-\lambda)]\eta_1\eta_2 - \lambda\phi}.$$

Solving (14) for ξ_t and writing the result with a general lag structure results

$$\xi_{t-i} = \frac{1}{\kappa - 1}\Delta s_{t+1-i}^e. \tag{15}$$

where $i = 1, \ldots, \infty$.

In order to write the dynamic complete model, we relax the hypothesis of a serially uncorrelated, ξ_t, and introduce dynamics into the reduced form by assuming serial correlation of the AR(2) type[6]

$$\xi_t = \theta_0 + \theta_1\xi_{t-1} + \theta_2\xi_{t-2} + \zeta_t, \tag{16}$$

where ζ_t is an independent white noise process. Using (15) into (16), generates

$$\xi_t = \theta_0 + \frac{\theta_1}{\kappa - 1}\Delta s_t^e + \frac{\theta_2}{\kappa - 1}\Delta s_{t-1}^e + \zeta_t, \tag{17}$$

and we finally substitute (17) into (14)

$$\Delta s_{t+1}^e = \theta_0(\kappa - 1) + \theta_1\Delta s_t^e + \theta_2\Delta s_{t-1}^e + (\kappa - 1)\zeta_t. \tag{18}$$

As will be shown, our tests reveal that serial correlation was eliminated when estimations were made using the fully dynamic model in (18).

[5] This draws on Fama (1984). For a comprehensive review on these relationships, see Engel (1996).
[6] Although this is a frequent assumption in the literature, it will be latter justified on empirical grounds.

3 Empirical Results

Exchange rates and Brazilian nominal interest rates were obtained from the Brazilian Central Bank, while United States of America (USA) nominal interest rates (three-month maturity Treasury Bill) are from the IFS/IMF (International Financial Statistics/International Monetary Fund). We used monthly interest rates and monthly expected exchange rate changes for the period that spans from 2001M11 until 2014M12.

An illustrative example might be helpful to understand the nature of our data. Consider, for instance, a random observation drawn from our sample, such as $2002M4$. Using approximate values, Brazilian nominal interest rate in $2002M4$ is 1.48% (19.28% annually compounded) while the USA treasury bill is 0.14 *per* month and 1.74% *per* year. End of period average exchange rate (i.e., the average between bid and ask exchange rates in the last working day of $2002M4$) is $2.3621R\$/US\$$, whereas the expected exchange rate for the last day of the following month, which was collected in the last working day of $2002M11$, is $2.583R\$/US\$$, giving an expected monthly depreciation rate of approximately 0.76%. The implied monthly (nominal) expected excess return is 0.58% or 7.18% on an annual compounded basis.

Descriptive statistics for the nid and Δs^e are presented in Table I. As can be seen, the average nominal interest rate differential is 0.95% *per* month while the expected change in the $R\$/US\$$ exchange rate is 0.13%, implying an average monthly excess return of 0.82% (10.29% yearly compounded, during the sample period).

Graph 1 plots the series for Δs^e_{t+1} in the right hand side axis and the nid_t on the left hand side axis. A feature of the data that stands out is the large drop in expected exchange rate changes from 2002 and 2003. The sharp depreciation of the Brazilian Real in that period, which explains this drop, was likely to be caused by a hike in risk due to uncertainty regarding the presidential election. The nominal interest rate differential seems to show an initial declining trend and a relatively stable path after 2005.

Table I: **Descriptive Statistics**

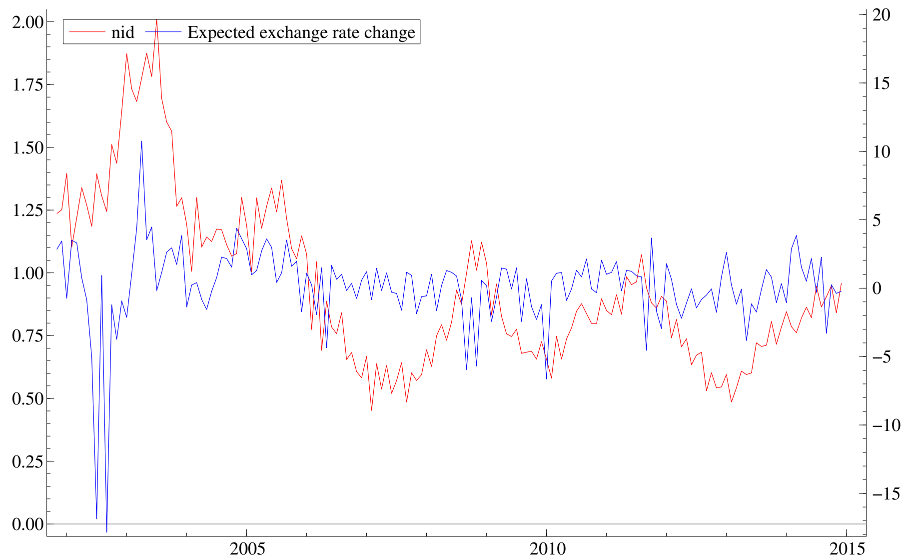
	Variable	
	nid	Δs^e_{t+1}
Mean	0.95	0.13
Standard Deviation	0.33	2.96
Maximum	2.01	10.75
Minimum	0.45	-17.83

Source: Brazilian Central Bank and IFS/IMF; 156 monthly observations.

For comparison purposes, we first estimated equation equation (11) using **observed exchange rate changes** in the place of expected changes. This is a common way in which the literature deals with the absence of observed data on expected exchange rate and it is frequently justified through a rational expectations hypothesis, under which a forecast error is added to ξ_t. Our OLS estimations reproduce the common finding, which is a negative slope parameter, equal to -0.23, associated to the nid. The estimated constant is equal to 0.38. Both parameters are not significant at reasonable confidence levels. Two Stage Least Squares (IV) estimation, using nid_{t-1} as an instrument, largely reduces the bias. The slope parameter increases to 0.41 but it is not significant - the associated t-value and corresponding probability are (0.28) and (0.77), respectively.

Estimations of equation (11) using survey data as a proxy for exchange rate expectations in the left hand side, as shown in Table II, present surprising results that are favorable to *ex ante* UIP. Although, the point estimate is above one, 1.25, if there is a negative bias of the type presented in our theoretical model, IV estimations would produce a higher coefficient. This is confirmed by our results, as the nid coefficient increases from 1.25 to 1.56 using nid_{t-1} as an instrument and two stage least squares, for example. As can be seen in Table II, addition of more lagged instruments reduce the bias almost monotonically. However, there is only one case in which we can reject the null of nid_t exogeneity. As shown by the Durbin/Wu/Hausman test, the p-value of the $\chi^2(1)$ statistics is 6.5% using the nid_{t-3} as instrument only (the latter is correlated with nid_t at a very small significance level). As a matter of fact, nid_{t-1} and nid_{t-2} will not be valid instruments under equation (16), which was assumed given that (18) better describes our data, according to the evidence presented below.

Figure 1: **Monthly changes in Δs^e_{t+1} and nid_t**



Source: Brazilian Central Bank and IMF/IFS.

Table II: **Estimations of Equation (11)**

The dependent variable is Δs^e_{t+1}						
	OLS	Instrumental Variables				
Instruments		nid_{t-1}	nid_{t-1} nid_{t-2}	nid_{t-1} nid_{t-2} nid_{t-3}	nid_{t-2} nid_{t-3}	nid_{t-3}
β_0	-1.05 (0.719)	-1.36 (0.678)	-1.36 (0.676)	-1.39 (0.679)	-1.42 (0.753)	-1.59 (0.765)
β_1	1.25 (0.718)	1.56 (0.853)	1.58 (0.836)	1.66 (0.843)	1.69 (0.941)	1.81 (0.940)
Durbin/Wu/Hausman test[1]		1.613 [0.204]	1.609 [0.204]	2.042 [0.152]	1.951 [0.162]	3.394 [0.065]

Notes: (1) Significance levels for $\chi^2(1)$ statistics of the Durbin/Wu/Hausman endogeneity tests (Durbin (1954), Wu (1973) and Hausman (1978)) are in brackets; (2) Standard errors of the respective estimated parameters are in parenthesis.

The earlier models' lack of predictive power is substantial (R^2 in the OLS estimation of Table II is 1.3%, for instance). This is due to the fact that the reduced form dynamically complete model better describes the nature of the observed data, rather than the structural equation. The estimated error of the *ex ante* UIP presented serial correlation. For example, we detected serial correlation of the second order in the OLS equation using survey data: the first autoregressive parameter is $\hat{\theta}_1 = 0.40$ and the second is $\hat{\theta}_2 = 0.12$ with associated p-values of 0.00 and 0.10, respectively, whereas the constant is not significant (similar results are found using nid_{t-3} as the single instrument). These results additionally justify testing equation (18). Evidence presented in Table III show that the estimated model does not have problems of serial correlation and the R^2 rises significantly, as expected. The remaining problems are related to heteroscedasticity, both unconditional and time conditional, and normality. We decided not to deal with these problems with the addition of time dummies or with the use of ARCH models, for example.

Table III: **OLS Estimation of Equation (18)**

The dependent variable is Δs^e_{t+1}			
	Coefficient	**Std Error**	**Significance**
Constant	0.02	0.2123	0.9105
ΔS^e_t	0.13	0.0736	0.0855
ΔS^e_{t-1}	0.41	0.0734	0.0000

Notes: $R^2 = 20.74\%$; Residual autocorrelation tests using lags from 1 to 2 and also from 1 to 6, produce statistics equal to $F(2,151) = 1.4797[0.2310]$ and $F(6,147) = 1.6202[0.1454]$, respectively, not rejecting the null of no autocorrelation.

4 Concluding Remarks

Our paper shows that UIP regressions can be substantially ameliorated using IV estimations and survey data on exchange rates. We also developed a simple macroeconomic model showing how a bias can arise when there is a simultaneous relation between agents and Central Banks.

Our model does not rely on the hypothesis of leaning against the wind and it is based on a monetary policy reaction function with smoothing. We showed that pass-through and a time-varying neutral real interest rate can explain the need of monetary authorities to react to UIP shocks. We were thus able to unveil the correlation between these shocks and nominal interest rates.

In summary, by showing that the simultaneity bias holds even when monetary authorities react to price changes, we complemented the work of McCallum (1994b) and Kugler (2000) also implying that Christensen (2000)'s results are not conclusive evidence against the simultaneity bias hypothesis.

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