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TIME HORIZON AND UNCOVERED INTEREST PARITY IN EMERGING ECONOMIES

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ABSTRACT

The aim of this study is to re-examine the well-known empirical puzzle of uncovered interest parity (UIP) for emerging market economies with different prediction time horizons. The empirical results obtained using dynamic panel and time series techniques for monthly data from January 1995 to December 2009 eventually show that the panel data estimates are more powerful than those obtained by applying individual time series estimations and the significant contribution of the exchange rate prediction horizons in determining the status of UIP. This finding reveals that at the longer time horizon, the model has better econometric specification and thus more predictive power for exchange rate movements compared to the shorter time period. The findings can also be a signalling of well-integrated currency markets and a reliable guide to international investors as well as for the orderly conduct of monetary authorities.

Keywords: uncovered interest parity, emerging markets, time series, panel co-integration

INTRODUCTION

Uncovered interest parity (UIP) is one of the oldest macroeconomic propositions and is still a building block of many international economic and finance theories. Contrary to widespread theoretical use of UIP, empirical tests of UIP reject the predicted relationship between interest rate differential and exchange rate changes. It is common to find an empirical result that shows the exchange rates of countries with high nominal interest rates tend to appreciate rather than depreciate in a short-term forecast horizon. Excellent reviews of the long-outstanding puzzle are provided by Engel (1996) and Chinn (2006). Some of the explanations offered for the rejection include the following: expectational errors (Mark & Wu, 1998; Kirikos, 2002), the presence of time-varying risk premia (Sarantis, 2006), inactivity-speculation zone (Cook, 2009; Paya, Peel, & Spiru, 2010) or policy behaviour (McCallum, 1994; Christensen, 2000; Chinn & Meredith, 2004).

Recently, some studies have attempted to find new ground for UIP by testing its validity at longer horizons. Fujii and Chinn (2001) show that the status of UIP could crucially depend on the long-term variables. Chinn and Meredith (2004) find that using longer maturity financial instruments (five- to ten-year bonds) substantially changes the sign of the interest rate coefficient from negative to positive, with three (out of six) currencies not rejecting the hypothesis that the slope coefficient is equal to 1. Augmenting McCallum's (1994) model, they argue that at short horizons, shocks in the exchange markets lead to monetary policy responses that result in a negative correlation between exchange rate changes and interest rate differentials. Contrary to short horizons, at longer time horizons, interest rates and exchange rates are both driven by macroeconomic "fundamentals" factors that result in a more consistent relationship with UIP.

However, Valkanov (2003) argues that using long-horizon regression could provide misleading statistical inferences compared to the short-horizon regression. Extra caution is required in long-horizon regression because of the overlapping sums of the original series (close to a unit root process) that might lead to *t*-statistics that do not converge to a well-defined asymptotic distribution. This situation may result in inconsistent ordinary least squared (OLS) estimators and inadequate measures for the coefficient of determination, R^2. Similar arguments are made by Kilian (1999). He employs bootstrap methods on monetary models to show that there is no significant increase in predictive power by using longer-horizon estimation methods.

The arguments used by Valkanov (2003) are no different than those made by Granger and Newbold (1974), and Phillips (1986). The analogy among them lies in finding a spurious correlation between persistent variables when they are statistically independent. All of these facts are related to the non-stationary behaviour that is usually exhibited by long-horizon variables.

All of the above-mentioned studies concentrate on developed and industrialised economies. Given the current status of liberalisation in emerging markets and their growing importance in global financial markets, in this paper, we re-examine UIP for emerging economies focusing on different time horizons to evaluate whether UIP holds or not. Furthermore, we use different based-currency for relative-country choice sensitivity as a means of checking robustness. Our main contributions to the literature are as follows. First, only very few studies dealt with UIP in emerging countries; among them are studies by Bansal and Dahlquist (2000), Flood and Rose (2001), Francis, Hasan and Hunter (2002) and Frankel and Poonawala (2010). This lack of studies on emerging countries exists because emerging markets were relatively closed until the mid-1980s. Previously, excessive constraints were imposed by local authorities either on capital movements or exchange rate changes, which makes the testing of UIP uninteresting. In this sense, we complement the existing literature on UIP because empirical work on emerging markets is still lacking; and second, the majority of studies considering emerging countries use short-term forecast horizons (k) in the regression of UIP models. For example, Bansal and Dahlquist (2000) use one- and three-month intervals, while Flood and Rose (2001), Francis et al. (2002), and Frankel and Poonawala (2010) use one-month horizons. Contrary to these papers, we extend the test of UIP by focusing on the different exchange and interest rate maturities from short to medium term, i.e., one-, three- and twelve-month horizons (k = 1, k = 3 and k = 12) using both dynamic time series and panel regression. Our findings for short-term forecast horizon confirm the earlier results for emerging economies (positive but still significantly different from 1). Interestingly, when we use longer forecast horizons (k = 12), the slope coefficients get closer to unity for most of the markets. As a robustness check, we further test the UIP hypothesis using different combinations of base countries.

The remainder of the paper is organised as follows. First, we briefly discuss the theory and recent evidence of UIP in emerging markets. Then, we describes the dataset used in the empirical analysis and the layout of the econometric procedures. Next, we discusses the estimation results. Finally, some concluding remarks are offered.

THEORY AND EVIDENCE IN EMERGING MARKETS

UIP states that the interest differential between two countries should equal the expected exchange rate changes. If the nominal interest rate in the foreign market is higher compared to the local market, it allows investors to borrow at the relatively low local rate and invest the proceeds at the higher foreign rate. Then, at the end of the k-th period, the foreign currency proceeds are converted back to local currency. The local currency is expected to appreciate just to reach an equilibrium point and cancel out the excess profit between these two markets. Ideally, this proposition holds true if the market satisfies the condition of no economic and/or political barriers (i.e., risk premium and political risk) between countries. In addition, the agents are assumed to be risk-neutral and behave rationally. Then, active arbitrage trading ensures that the UIP hypothesis holds. On the contrary, if this phenomenon does not hold, there is ample room for trading manipulation, which eventually leads to market inefficiency.

The above explanation is one specification of UIP, which can be expressed in the following equation:

$$\Delta s_{t,k} = \alpha + \beta(i_{t,k} - i^*_{t,k}) + \varepsilon_{t,k} \tag{1}$$

where $\Delta s_{t,k}$ is the change of the domestic exchange rate over time period *k*, $(i_{t,k} - i^*_{t,k})$ is the interest rate differential between domestic and foreign markets for maturity in *k* periods, subscript *t* represents time, and $\varepsilon_{t,k}$ is an error term. Given that markets are efficient with regard to arbitrage activities and neither political nor economic barriers exist between markets, the estimated parameters of α and β should not be statistically different from 0 and 1, respectively, and the error term should be white noise. The failure of any hypothesis from which the model is derived indicates the presence of a time-varying risk premium.

Testing of UIP in emerging markets is still relatively lacking. This deficiency may be for at least two reasons. The first is the relatively fixed exchange rate regimes and extensive controls on the economy in some of these markets until the mid-1980s and early 1990s. These restrictions violate the theoretical framework of UIP and may cause the "peso problem" in its empirical testing (Krasker, 1980). In this study, we try to avoid this problem by dropping countries with excessive capital control and adopting hard peg exchange rate regimes. We thus consider only countries that have a free capital account and a relatively floating exchange rate regime, which allows the exchange rate to fluctuate, i.e., from a band to a free-floating regime. Recent literature has found the difficulty in establishing whether a declared flexible or fixed exchange rate regime is in fact just *de jure* or also *de facto* (Reinhart & Rogoff, 2004).

Second, the dearth of *ex ante* exchange rate datasets. Strictly speaking, UIP is an *ex ante* concept defined by expectations rather than *ex post* realised depreciation rates. To avoid this problem, many researchers, such as Francis et al. (2002) and Cheung, Chinn and Fujii (2005, 2006), carry out an investigation of Rational Uncovered Interest Parity (RUIP) in emerging markets by assuming rational expectations and using an *ex post* instead of an *ex ante* series. Bansal and Dahlquist (2000) use a latent factor model for both cross-sectional and time series data from 12 emerging economies to show that UIP performs better in emerging economies compared to developed economies. Their findings indicate that the deviation from UIP occurs only in two specific scenarios; the first is when the U.S. interest rate exceeds the foreign interest rate, and the second is if the foreign interest rate is higher than the local rate. Bansal and Dahlquist also find that country-specific attributes, such as *per capita* income, inflation, volatility, country risk rating and nominal interest rate, are important in explaining the deviation from the UIP hypothesis.

Francis et al. (2002) further investigate the empirical puzzle of UIP for 9 developing countries (Chile, Columbia, Mexico, India, Korea, Malaysia, Pakistan, Thailand and Turkey) in pre- and post-liberalisation eras using a multi-factor conditional asset-pricing model estimated in a multivariate GARCH framework. This research confirms that the deviation from UIP prevails in most of the emerging countries and that the phenomenon is country-specific in nature.

Using the one-month forward exchange rate, Frankel and Poonawala (2010) test the unbiasedness hypothesis for fourteen emerging countries from 1996 to 2004. The results from the individual market time-series regressions are mixed. Eight markets experienced a positive estimated forward-discount coefficient, β (although smaller than unity), and the remaining were negative and statistically insignificant. They also find a positive slope for β by pooling together the emerging countries.

To summarise, the evidence against the UIP puzzle in the post-liberalisation era in emerging economies is not as severe as was commonly thought in the pre-liberalisation period. However, the evidence is still far from conclusive, and it is country-specific in nature.

DATA AND ECONOMETRIC SPECIFICATION

Data Description

In this study, the UIP hypothesis was tested using monthly data of exchange rate changes and interest rate differentials spanning from January 1995 to December 2009 for 15 emerging markets with the U.S. as a base country (hereafter, we call this a U.S.-base model). The countries included were four Latin-American emerging markets (Brazil, Chile, Mexico and Venezuela), four Asian emerging markets (Indonesia, Korea, the Philippines and Thailand), five European emerging markets (Hungary, Poland, Portugal, Romania and Russia), one Middle-Eastern emerging market (Israel) and an African emerging market (Morocco). These emerging countries were selected based on the capital account openness and an exchange rate regime that at least allows for large exchange rate movement throughout the sample period. However, countries with hard peg exchange rate regimes in some of the sample period, such as Malaysia (1998 to 2005) and Argentina (1991 to 2001), or capital control regimes, such as India, were omitted from the dataset.

The interest rates used are the 1-month, 3-month and 12-month deposit rate, inter-bank rate or Treasury bill rate of monthly frequency. All interest rate series were downloaded from Datastream. The monthly exchange rate series were extracted from the International Financial Statistics (IFS) and expressed in terms of U.S. dollars per unit of emerging market currency. Details of the data set used in the analysis are presented in Table 1.

Table 1
Monthly data specification for emerging countries from January 1995 to December 2009

	Interest rate	Time horizon	Period	Base country
Brazil	Deposit	1-month	Jan. '95 – Dec. '09	U.S., Germany
Chile	Deposit	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Japan
Mexico	Deposit	1-, 3-, 12-month	Jan. '96 – Dec. '09	U.S., Japan
Venezuela	Deposit	1-, 3-month	Dec. '96 – Dec. '09	U.S., Germany
Indonesia	Deposit	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Japan
Korea	Deposit	3-, 12-month	Jan. '95 – Dec. '09	U.S., Japan
Philippines	Deposit	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Japan
Thailand	Deposit	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Japan
Israel	T-Bill	3-, 12-month	Jan. '95 – Dec. '09	U.S., Germany
Morocco	Deposit	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Germany
Hungary	Interbank	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Germany
Poland	Interbank	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Germany
Portugal	Interbank	1-, 3-month	Jan. '95 – Dec. '09	U.S., Germany
Romania	Interbank	1-, 3-, 12-month	Jan. '95 – Dec. '09	U.S., Germany
Russia	Interbank	1-, 3-month	Jan. '95 – Dec. '09	U.S., Germany

Note: Data for nominal interest rates are collected from Datastream. The selection of relative country is base on the first two largest trading partners with respective emerging economies in direction of trade (DoT) statistics.

Econometric Methodology

The empirical analysis of Equation 1 (refer page 6) is carried out by developing the following basic steps for the three different models, i.e., U.S.-base model, Japan-base model and German-base model. The name of the model is chosen depending on the relative country used in the exchange rate arrangements.

Time Series Analysis

For preliminary analysis, we implemented unit root tests using the Augmented Dickey Fuller (ADF) test in level and first difference of the series covering various time-lag terms. All the series including changes in exchange rate or difference of local and foreign interest rate (as in Equations 1, 2 and 3) needed to be thoroughly investigated for their stationarity level because all these series have different persistency properties due to the different forecasting horizon k. The optimal lag was chosen using the Akaike Information Criteria (AIC) specification. The results of the test applied to the series in level indicate that we did not reject the null hypothesis of a unit root for interest rate differential at all horizons and for all markets, except for the 1-month and 3-month maturities of Romania and Russia. In the case of exchange rate changes, we only failed to reject the null of the unit root at the 12-month horizon for all countries. The first difference series were stationary. In general, the results show that all interest rate series are I(1), while exchange rates are I(0) for 1-month and 3-month horizons and I(1) for the 12-month horizon. Table 2 provides a summary of the ADF unit root tests.

Due to the stationarity property of 1-month and 3-month horizons for dependent variables (exchange rate changes), and becomes non-stationary at k = 12, we estimated UIP using two different procedures. First, we used the standard OLS method for k = 1 and k = 3 with additional dummy variables to capture the crises that affected some of the countries during the sample period, i.e., the Asian financial crisis in 1997 and the Russian crisis in 1998. The Newey-West robust standard errors were used to give consistent covariance matrices in the presence of both serial correlation and heteroskedasticity.

Table 2
Summary of unit root properties of exchange rate movement and interest rate differential using ADF unit root tests for data from January 1995 to December 2009

Country			Exchange rate			Interest rate
	1-m	3-m	12-m	1-m	3-m	12-m
Brazil	$I(0)$	–	–	$I(1)$	–	–
Chile	$I(0)$	$I(0)$	–	$I(1)$	$I(1)$	–
Mexico	$I(0)$	$I(0)$	$I(1)$	$I(1)$	$I(1)$	$I(1)$
Venezuela	$I(0)$	$I(0)$	–	$I(1)$	$I(1)$	–
Indonesia	$I(0)$	$I(0)$	$I(1)$	$I(1)$	$I(1)$	$I(1)$
Korea	–	$I(0)$	$I(1)$	–	$I(1)$	$I(1)$
Philippines	$I(0)$	$I(0)$	$I(1)$	$I(1)$	$I(1)$	$I(1)$
Thailand	$I(0)$	$I(0)$	$I(1)$	$I(1)$	$I(1)$	$I(1)$
Israel	–	$I(0)$	$I(1)$	–	$I(1)$	$I(1)$
Morocco	$I(0)$	$I(0)$	$I(1)$	$I(1)$	$I(1)$	$I(1)$
Hungary	$I(0)$	$I(0)$	$I(1)$	$I(1)$	$I(1)$	$I(1)$
Poland	$I(0)$	$I(0)$	$I(1)$	$I(1)$	$I(1)$	$I(1)$
Portugal	$I(0)$	$I(0)$	–	$I(1)$	$I(1)$	–
Romania	$I(0)$	$I(0)$	$I(1)$	$I(0)$	$I(0)$	$I(1)$
Russia	$I(0)$	$I(0)$	–	$I(0)$	$I(0)$	–

Note: $I(0)$ refers to stationary at level form and $I(1)$ refers to stationary at first difference. We used 1% and 5% critical value that was provided by MacKinnon (1996) to test the significance level. The lag length has been selected based on AIC to ensure white noise residual. – indicates non availability of series.

Second, due to the persistency problem in dependent and independent variables for k = 12, we employed Stock and Watson's (1993) Dynamic OLS (DOLS) to estimate the long-run parameters of UIP for k = 12. The DOLS procedure basically involves regressing any co-integrated $I(1)$ variables on other $I(1)$ variables, any $I(0)$ variables and leads and lags of the first differences of any $I(1)$ variables. The procedure can be represented in the following econometric specification:

$$\Delta s_{t,k} = \alpha + \beta_D (i_{t,k} - i^*_{t,k}) + \sum_{q=-q_1}^{q_2} \delta_q \Delta (i_{t-q,k} - i^*_{t-q,k}) + \varepsilon_{t,k} \tag{2}$$

where β_D is the Stock-Watson DOLS parameter which estimates the long-run parameters with the interest rate differential appearing in level. q is the optimum number of lead and lag terms included in the estimation to provide an efficient estimator of the co-integrating coefficient. We also use the heteroskedasticity consistent covariance proposed by Newey and West (1987) to avoid the problem of whether or not the regression errors are heteroskedastic and autocorrelated.

Panel Analysis

Panel data estimates are more powerful than those obtained by applying individual time series estimations, especially in short-span data sets. Levin, Lin and Chu (2002) argue that panel analysis will eventually increase the power of the test and minimise the problem of statistical inferences.

The empirical investigation test procedure was conducted using the following steps. First, we investigated the unit root properties for each cross-section using methodology proposed by Levin, Lin and Chu (2002), which will be referred to as LLC hereafter, and Im, Pesaran and Shin (2003), which will be referred to as IPS hereafter. We tested the null of the unit root by comparing the IPS *w*-statistics and LLC t^*-statistics to 95% critical values. These two techniques are robust over the problems of homogeneity and heterogeneity across units on the lagged variable.

Second, for k = 1 and k = 3, in which exchange rate depreciation and interest rate differential are stationary, we employed the standard panel OLS techniques to Equation 1 with and without fixed effect.

Alternatively, for k = 12, in which both series are persistent and non-stationary, we utilised two types of the heterogeneous panel co-integration test developed by Pedroni (1999; 2004) and Kao (1999). Basically, both Pedroni and Kao extend the Engel-Granger two-step residual-based co-integration framework to tests involving panel data for the following equation:

$$\Delta s_{jt,k} = \sum_{q=1}^{Q} \beta_{jq} \Delta (i_{jt-q,k} - i^*_{jt-q,k}) + \mu_{jt,k} \tag{3}$$

where subscript j is an individual emerging economy and the Q is the AIC optimal lag number. In this study, specifically we considered two types of the heterogeneous panel co-integration test developed by Pedroni (1999; 2004), which allows different individual effects across cross-sectional interdependency. The first type of test includes the panel rho (ρ), panel non-parametric (*PP*) and panel parametric (ADF) statistics. The panel parametric statistics are similar to the single-equation ADF-test, and panel non-parametric statistics are analogous to the Phillips and Perron (1988) test. The second type of test proposed by Pedroni (1999; 2004) is comparable to the group mean panel tests of Im, Pesaran and Shin (2003). Pedroni argues that both types of test are appropriate for testing the null of co-integration in bivariate panel models with heterogeneous dynamic, fixed effects and heterogeneous co-integrating slope coefficients. Further, Pedroni claims that this method also will take into account the off-diagonal terms in the residual long-run covariance and the effect of spurious regression in the heterogeneous panel.

Further, we considered the panel co-integration tests of Kao (1999). The Kao test follows the same basic approach as the Pedroni tests but specifies cross-section specific intercepts and the homogeneous coefficient on the first stage regressors. The limiting distribution of the residual-based co-integration tests using the *DF* test and *ADF*. Under the null of no co-integration, Kao shows that all the DF_ρ, DF_t, DF_ρ^*, DF_t^*, and *ADF* test statistics are converged to a standard normal asymptotic distribution.

If there was evidence of co-integration, we further estimated the co-integration coefficients for the panel using bias-corrected ordinary least squares (bias-corrected-OLS), fully modified ordinary least squares (FM-OLS) and dynamic ordinary least squares (DOLS) under the homogenous covariance structure proposed by Kao and Chiang (2000). We used these three different methods to avoid and compare any estimation bias at longer horizons. The Kao's DOLS specification can be represented as follows:

$$\Delta s_{jt,k} = \alpha_j + \beta_D (i_{jt,k} - i^*_{jt,k}) + \sum_{q=-q_1}^{q_2} \delta_{jq} \Delta (i_{jt-q,k} - i^*_{jt-q,k}) + \mu_{jt,k} \tag{4}$$

The parameter α_j is the member-specific intercept or a fixed-effect parameter to cater for omitted variables that differ between markets but are constant over time. β_D is the DOLS long-run parameter estimate, and q is the number of lead and lag terms to correct the nuisance parameter to obtain coefficient estimates with nice limiting distribution properties as described by Kao and Chiang (2000).

EMPIRICAL RESULTS

Time Series Analysis

Table 3 and Table 4 depict the results of the country-by-country standard OLS coefficient (β_O) for k = 1 and k = 3, respectively, while Table 5 presents the DOLS (β_D) for k = 12. Because both exchange rate and interest rate differentials for k = 12 are of first differenced stationary series I(1), it is necessary to check whether these two series are co-integrated to ensure the β_D estimates are efficient.

The last column of Table 5 under the ADF heading shows the bivariate residual-based two-step co-integration test for k = 12 using the ADF technique. All ADF statistics are much smaller than the critical values, which leads to the conclusion that we rejected the null hypothesis of a unit root for all estimated residuals for all emerging market models irrespective of their relative countries (the U.S., Japan or Germany). This finding confirms that the exchange rate changes and interest rate differentials in these markets are co-integrated. Therefore, the Stock-Watson parameter estimates of the long-run parameter (β_D) are valid and not spurious. This time-series model (Equation 2) was estimated, including up to q = ± 3 leads and lags, without altering the results to any significant degree.

The striking result of the estimated coefficient for U.S.-based regression, β (inclusive of both β_O and β_D), is that at longer horizons (higher k), the UIP regression tends to produce estimates that are positive and not significantly different from unity. In Table 5 Panel A, when k = 12, nine β_D estimates are positive and statistically significant compared to only five and two for k = 3 in Table 4 and k = 1 in Table 3, respectively. Furthermore, five β_D estimates out of nine are statistically not different from unity. The results discussed above are robust because the same pattern of results is also reported for the UIP regression under the Japan and Germany models (Panel B and Panel C of Table 3, Table 4 and Table 5, respectively).

Table 3

Ordinary Least Square (OLS) regression for individual emerging market for k = 1 from 1995 to 2009 for equation $\Delta s_{t,k} = \alpha + \beta_O (i_{t,k} - i^*_{t,k}) + \varepsilon_{t,k}$

Country	α	SE (α)	β_O	SE (β_O)	$\beta_O = 1$	$\overline{R}^2$
A: U.S.						
Brazil ♣	0.919	(1.368)	0.070	(0.061)	0.000	0.089
Chile	0.663	(0.699)	–0.259	(0.170)	0.000	0.015
Mexico	0.236	(0.388)	0.046	(0.034)	0.000	0.008
Venezuela	0.654	(1.056)	0.142	(0.093)	0.000	0.012
Indonesia♣	–4.255*	(1.891)	–0.464	(0.236)	0.000	0.191
Korea♣	–	–	–	–	–	–
Philippines♣	–0.300	(0.457)	0.109	(0.194)	0.000	0.054
Thailand♣	1.340	(1.209)	–1.981	(1.322)	0.026	0.002
Israel	–	–	–	–	–	–
Morocco	–1.59***	(0.375)	–0.367***	(0.097)	0.000	0.083

(*continued*)

Table 3 (*continued*)

Country	α	SE (α)	β_O	SE (β_O)	$\beta_O = 1$	$\overline{R}^2$
Hungary	–15.90**	(2.788)	–0.551**	(0.101)	0.000	0.121
Poland	0.767	(0.598)	0.086	(0.043)	0.000	0.015
Portugal	–0.155	(0.248)	0.233*	(0.138)	0.000	0.030
Romania	1.681**	(0.706)	0.090***	(0.019)	0.000	0.552
Russia♣	0.284	(0.439)	0.046	(0.038)	0.000	0.246
B: Japan						
Chile	–0.246	(0.647)	–0.393	(1.369)	0.000	0.027
Mexico	0.034	(0.757)	0.022	(0.042)	0.000	0.014
Indonesia♣	–4.746	(2.649)	–0.366	(0.232)	0.000	0.110
Korea♣	–	–	–	–	–	–
Philippines♣	–0.146	(1.652)	0.049	(0.284)	0.000	0.072
Thailand♣	–0.105	(0.582)	0.115	(0.277)	0.001	0.019
C: Germany						
Brazil	0.010	(1.362)	0.017	(0.053)	0.000	0.141
Venezuela	0.144	(1.288)	0.101	(0.100)	0.000	0.023
Israel	–	–	–	–	–	–
Morocco	0.950	(1.704)	0.214	(0.330)	0.019	0.020
Hungary	–11.653	(11.004)	–0.385	(0.484)	0.000	0.028
Poland	1.331	(0.845)	0.150***	(0.060)	0.000	0.043
Portugal	0.136	(0.560)	0.386*	(0.231)	0.009	0.060
Romania	2.175**	(0.920)	0.101***	(0.022)	0.000	0.409
Russia♣	0.293	(0.595)	0.048	(0.033)	0.000	0.221

Note: $SE(\bullet)$ is Newey-West Standard Errors. $\beta_O = 1$ refers to *p-value* of the *F*-statistic. ***, ** and * indicate significant at 1%, 5% and 10% respectively.♣ Financial crisis dummy has been considered in the regression. – indicates non availability of dataset.

Table 4
Ordinary Least Square (OLS) regression for individual emerging market for k = 3 from 1995 to 2009 for equation $\Delta s_{t,k} = \alpha + \beta_O (i_{t,k} - i^*_{t,k}) + \varepsilon_{t,k}$

Country	α	SE (α)	β_O	SE (β_O)	$\beta_O = 1$	$\overline{R}^2$
A: U.S.						
Brazil ♣						
Chile	1.327	(1.899)	–0.519	(0.442)	0.000	0.107

(*continued*)

Table 4 (*continued*)

Country	α	SE (α)	β_O	SE (β_O)	$\beta_O = 1$	$\overline{R}^2$
Mexico	0.203	(0.881)	0.123*	(0.059)	0.000	0.277
Venezuela	–0.413	(2.416)	0.251	(0.191)	0.000	0.112
Indonesia♣	–7.252*	(3.010)	0.716**	(0.343)	0.000	0.452
Korea♣	1.073	(1.461)	0.445	(0.389)	0.156	0.451
Philippines♣	1.228	(2.431)	0.588	(0.499)	0.411	0.221
Thailand♣	1.355	(2.822)	–3.086	(2.571)	0.000	0.150
Israel	–0.429	(0.671)	0.145	(0.100)	0.000	0.118
Morocco	–1.35***	(0.475)	–1.260***	(0.263)	0.000	0.189
Hungary	–50.56**	(6.494)	–1.740**	(0.236)	0.000	0.398
Poland	1.648	(1.644)	0.239*	(0.121)	0.000	0.044
Portugal	–0.165	(0.635)	0.833**	(0.364)	0.000	0.047
Romania	5.382**	(2.075)	0.288***	(0.057)	0.000	0.674
Russia♣	1.404	(1.382)	0.132	(0.088)	0.000	0.454
B: Japan						
Chile	–0.906	(1.162)	–2.499	(3.823)	0.000	0.019
Mexico	–0.374	(2.171)	0.034	(0.117)	0.000	0.033
Indonesia♣	–7.454	(4.149)	–0.555	(0.337)	0.000	0.410
Korea♣	1.160	(2.169)	0.2333	(0.325)	0.020	0.353
Philippines♣	0.192	(2.865)	0.203	(0.331)	0.018	0.119
Thailand♣	–0.228	(1.541)	0.070	(0.588)	0.000	0.057
C: Germany						
Brazil	–	–	–	–	–	–
Venezuela	–2.728	(2.760)	0.088	(0.190)	0.000	0.042
Israel	1.580	(1.808)	0.490*	(0.199)	0.012	0.120
Morocco	1.505	(2.191)	0.911	(0.879)	0.919	0.023
Hungary	–57.838	(37.030)	–1.956	(1.282)	0.000	0.059
Poland	3.729	(2.357)	0.446**	(0.174)	0.002	0.136
Portugal	0.349	(1.586)	0.677	(0.782)	0.681	0.023
Romania	7.23***	(2.688)	0.333***	(0.063)	0.000	0.569
Russia♣	1.204	(1.761)	0.122	(0.080)	0.000	0.438

Note: SE is Newey-West Standard Errors. $\beta_O = 1$ refers to *p-value* of the *F*-statistic. ***, ** and * indicate significant at 1%, 5% and 10% respectively.♣ Financial crisis dummy has been considered in the regression. – indicates non availability of dataset.

Table 5
Stock-Watson dynamic ordinary least square (DOLS) regression for k = 12 for individual emerging market from 1995 to 2009 for equation

$$\Delta s_{t,k} = \alpha + \beta_D (i_{t,k} - i^*_{t,k}) + \sum_{q=-q_1}^{q=q_2} \delta_i \Delta (i_{t-n,k} - i^*_{t-q}) + \varepsilon_{t,k}$$

Country	α	SE (α)	β_O	SE (β_O)	$\beta_O = 1$	$\overline{R}^2$	ADF
A: U.S.							
Chile	–3.219	(1.611)	2.980***	(0.883)	0.028	0.189	–3.191**
Mexico	2.697	(2.493)	0.647***	(0.140)	0.013	0.339	–2.965**
Indonesia♣	2.390	(4.454)	0.859**	(0.412)	0.733	0.771	–3.080**
Korea♣	9.506**	(3.557)	2.587**	(0.691)	0.023	0.614	–3.91***
Philippines♣	7.700*	(4.603)	1.764***	(0.705)	0.280	0.702	–2.963**
Thailand♣	–3.973**	(1.855)	2.584**	(1.109)	0.155	0.431	–3.368**
Israel	0.317	(1.700)	0.897**	(0.276)	0.711	0.337	–2.747**
Morocco	–5.42***	(1.113)	–2.74***	(0.816)	0.000	0.193	–2.714**
Hungary	–219.1**	(16.969)	–7.495**	(0.604)	0.000	0.857	–4.51***
Poland	11.55***	(2.497)	1.575***	(0.191)	0.003	0.603	–3.038**
Romania	17.774***	(2.569)	1.167***	(0.106)	0.118	0.724	–5.67***
B: Japan							
Chile	87.141	(47.377)	14.974	(7.542)	0.039	0.123	–2.017**
Mexico	–5.019	(5.075)	–0.054	(0.257)	0.000	0.030	–2.894**
Indonesia♣	5.210	(6.378)	0.784	(0.440)	0.625	0.680	–3.185**
Korea♣	2.608	(4.400)	0.146	(0.639)	0.184	0.456	–3.82***
Philippines♣	8.234	(7.418)	1.373*	(0.723)	0.607	0.130	–2.70***
Thailand♣	–0.994	(3.323)	1.010	(1.041)	0.991	0.181	–3.07**
C: Germany							
Israel	6.758	(4.685)	1.849**	(0.550)	0.126	0.183	–2.651**
Morocco	6.057	(7.628)	2.590	(2.092)	0.448	0.055	–2.629**
Hungary	–550**	(77.530)	–18.59**	(2.643)	0.000	0.531	–2.008**
Poland	24.8***	(3.133)	2.88***	(0.220)	0.000	0.794	–3.76**
Romania	29.5***	(3.749)	1.462***	(0.124)	0.003	0.741	–5.19**

Note: SE is Newey-West Standard Errors. $\beta = 1$ refers to *p-value* of the *F*-statistic. ***, ** and * indicate significant at 1%, 5% and 10% respectively.♣ Financial crisis dummy has been considered in the regression. ADF is unit root test for $\varepsilon_{t,k}$ of Equation 1 and test using the critical value from MacKinnon (1991).

These results are consistent with previous empirical UIP testing in emerging markets in which emerging markets' regression generally produces more favourable results compared to developed markets, as documented in Bansal and Dahlquist (2000), Madarassy and Chinn (2002), and Frankel and Poonawala (2010). However, after considering a longer-term forecast horizon, the phenomenon of appreciation in the exchange rate in high nominal interest countries, such as Morocco and Hungary, remains an empirical puzzle for the UIP framework.

Panel Analysis

Prior to testing for panel regression and co-integration, LLC and IPS panel unit root tests were carried out, and the results are presented in Table 6. The results clearly show that the IPS *w*-statistics and LLC t^*-statistics reject the null hypothesis of a unit root at 5% only at first difference for both k = 12 for exchange rate and interest rate differentials. In contrast, the results for k = 1 and k = 3 have a mixed combination of 'reject' and 'fail to reject' IPS *w*-statistics and LLC t^*-statistics for interest rate differentials at level form. This finding is true whether or not we allow for a deterministic trend to appear in the unit root test specification. Generally, the results are consistent with individual series, in which both variables are differenced stationary $I(1)$ at k = 12, while for k = 1 and k = 3, exchange rates are stationary at level, but interest rates are only stationary at first difference. For k = 12, we need to further confirm whether these two $I(1)$ variables are co-integrated to establish an efficient long-run relationship. Table 7 shows the bivariate panel co-integration test proposed by Kao (1999) and Pedroni (1999; 2004). All test statistics for Kao (1999), i.e., DF_ρ, DF_t, DF_ρ^*, DF_t^* and ADF, reject the null of no co-integration at the 1% significance level for all models. For Pedroni test statistics, as indicated by the panel non-parametric (Z_{pp}-statistics) and parametric ($\tilde{Z}_t$-statistics) as well as by their group statistics, the null is rejected at the 1% level of significance for the U.S. model. The Japan and Germany models are also supports for co-integration between these two variables to show the robustness of the results. Both the Pedroni and Kao panel co-integration tests are consistent and confirm that at longer maturity periods taken as a group, exchange rate and interest rate differentials are co-integrated, and this finding could be an indication of the existence of the UIP.

Table 6
Panel unit root test for U.S., Japan and Germany models from January 1995 to December 2009

Variables			U.S.		Japan		Germany	
A: Level			LLC	IPS	LLC	IPS	LLC	IPS
	$k = 1$	No trend	–35.8**	–33.11**	–26.9**	–27.2**	–16.4**	–15.3**
		Trend	–41.39**	–34.78**	–29.9**	–28.3**	–24.2**	–19.5**
Exchange rate	$k = 3$	No trend	–5.88**	–8.46**	–3.55**	–8.46**	–3.22**	–5.60**
		Trend	–6.95**	–8.34**	–3.04**	–7.39**	–5.51**	–6.88**
	$k = 12$	No trend	0.30	–0.86(13)	–0.33	–3.51**	–0.73	–0.86
		Trend	2.23	–1.32(2)	0.34	–1.28(1)	–1.12	–0.45
	$k = 1$	No trend	–2.71**	–4.26**	–0.91	–2.41**	–3.72**	–4.08**
		Trend	–1.24	–3.50**	–1.07	–2.47**	–3.28**	–4.15**
Interest rate	$k = 3$	No trend	–1.30	–2.29*	–1.06	–2.14*	–4.29**	–3.37**
		Trend	–0.23	–2.42**	–0.24	–2.70**	–3.08**	–3.46**
	$k = 12$	No trend	–0.15	–1.00	–1.12	–1.05	–1.03	–1.23
		Trend	–0.49	–0.88	–1.26	–1.50	–0.80	–1.02

(*continued*)

Table 6 (*continued*)

Variables			U.S.		Japan		Germany	
B: First difference			LLC	IPS	LLC	IPS	LLC	IPS
	$k = 1$	No trend	–	–	–	–	–	–
		Trend	–	–	–	–	–	–
Exchange rate	$k = 3$	No trend	–	–	–	–	–	–
		Trend	–	–	–	–	–	–
	$k = 12$	No trend	–33.28**	30.82**	–31.4**	–	–16.8**	–15.2**
		Trend	–36.97**	31.31**	–34.9**	–30.9**	–18.8**	–14.9**
B: First difference			LLC	IPS	LLC	IPS	LLC	IPS
	$k = 1$	No trend	–	–	–18.3**	–	–	–
		Trend	–37.48**	–	–20.6**	–	–	–
				–		–	–	–
Interest rate	$k = 3$	No trend	–32.31**	–	–26.0**	–	–	–
		Trend	–36.27**	–	–29.3**	–	–	–
	$k = 12$	No trend	–28.78**	–26.79**	–24.9**	–21.63**	–19.97**	–17.50**
		Trend	–32.13**	–27.03**	–27.6**	–21.90**	–22.35**	–17.66**

Note: IPS is *w*-statistic from Im et al. (2003) and LLC is t^*-statistic from Levin et al. (2002). * and ** indicate significant at 5% and 1% level.

The analysis is pursued, therefore, by estimating the co-integrating coefficient using panel bias corrected OLS, FMOLS and DOLS under the heterogeneous covariance structure proposed by Kao and Chiang (2000) for $k = 12$ and standard panel OLS for $k = 1$ and $k = 3$. The results for estimated coefficients with their t-statistics in parentheses are presented in Tables 8 and 9, respectively. One main feature of the results is that the estimated interest rate differential coefficient has the correct sign as predicted by the hypothesis (positive) and is getting closer to unity at longer time horizons for all models. For instance, for the U.S.-base model, as maturity (k) increases from 1 to 3 and then to 12, β increases from 0.05 to 0.20 and 0.641 (for DOLS or 1.064 for bias-corrected and 0.681 for FMOLS), respectively. Statistically, the β from the DOLS estimate is superior to the other two estimates (Kao & Chiang, 2000). The other two models (Japan-base and German-base) produce the same pattern of β as k increases from 1 to 3 and then to 12. This finding is more favourable than the existing literature in which Bansal and Dahlquist (2000) find the pool coefficient on interest rate differential for developing markets for 3-month maturity to be 0.19. However, Bansal and Dahlquist (2000) do not proceed further with longer maturity periods to show the pattern of β as k increases. Our finding, which is new for emerging markets, is quite similar to Chinn and Meredith (2005) who find the panel coefficient on interest rate differential to be around 0.674 at 5-year maturity for developed markets. This finding indicates that, consistent with the individual series regression, the estimated coefficient of interest rate differential in emerging markets is positive and it is converging to unity at longer horizons of k.

Table 7
Panel cointegration for U.S., Japan and Germany models from January 1995 to December 2009

	U.S.	Japan	Germany
	$N = 11$	$N = 6$	$N = 5$
A: Kao (1999)			
DF_ρ	–7.539**	–11.35**	–6.03**
DF_t	–3.997**	–5.49**	–3.37**
DF^*_ρ	–17.42**	–24.03**	–14.95**
DF^*_t	–3.739**	–4.62**	–3.31**
ADF	–5.23**	–4.85**	–4.82**

(*continued*)

Table 7 (*continued*)

	U.S.	Japan	Germany
	$N = 11$	$N = 6$	$N = 5$
B: Pedroni (1999; 2004)			
Intercept and no trend			
Z_ρ	–1.504	–0.843	–2.099*
Z_{pp}	–3.524**	–2.257*	–1.521
Z_t	–5.286**	–2.955**	–4.171**
$\tilde{Z}_\rho$	–0.663	–0.432	–0.897
$\tilde{Z}_{pp}$	–3.778**	–2.548	–0.688
$\tilde{Z}_t$	–5.318**	–2.830**	–2.905**
Intercept and trend			
Z_ρ	–0.849	–0.194	–0.982
Z_{pp}	–2.061*	–2.109*	–0.448
Z_t	–2.996**	–2.444*	–2.048*
$\tilde{Z}_\rho$	–0.572	–0.477	–0.110
$\tilde{Z}_{pp}$	–2.817**	–1.811	–0.423
$\tilde{Z}_t$	–4.147**	–2.181*	–0.829

Note: Cointegration test statistics are calculated through the residuals from the panel OLS estimation. * and ** indicate significant at 5% and 1% level respectively.

Table 8
Panel OLS regression of UIP for emerging markets from 1995 to 2009

	U.S.		Japan		Germany	
	k = 1	*k = 3*	*k = 1*	*k = 3*	*k = 1*	*k = 3*
A: Fixed Effect						
α	–0.054	0.315	–0.313	–0.414	0.283	1.548**
	(0.160)	(0.274)	(0.387)	(0.706)	(0.244)	(0.420)
β	0.050**	0.200**	0.010	0.088	0.060**	0.232**
$se(\beta)$	(0.008)	(0.016)	(0.036)	(0.069)	(0.009	0.017)

(*continued*)

Table 8 (*continued*)

	U.S.		Japan		Germany	
	k = 1	*k = 3*	*k = 1*	*k = 3*	*k = 1*	*k = 3*
A: Fixed Effect						
$\beta = 1$	0.000	0.000	0.000	0.000	0.000	0.000
$\bar{R}^2$	0.047	0.145	0.007	0.012	0.071	0.225
Obs	1662	1759	643	760	1019	999
NoID	13	14	5	6	8	8
B: No Fixed Effect						
α	–0.125	0.046	–0.229	–0.324	0.206	1.198*
$se(\alpha)$	(0.145)	(0.358)	(0.327)	(0.575)	(0.235)	0.575
β	0.045**	0.176	0.020	0.098*	0.056**	0.214**
$se(\beta)$	(0.006)	(0.013)	(0.026)	(0.049)	(0.007)	0.016
$\beta = 1$	0.000	0.000	0.000	0.000	0.000	0.000
$\bar{R}^2$	0.045	0.098	0.003	0.014	0.064	0.165
Obs	1662	1759	1662	1759	+	1759
NoID	13	14	13	14	13	14

Note: Panel regression of [$\Delta s_{it,k} = \alpha + \beta(i_{it,k} - i^*_{it,k}) + \varepsilon_{it,k}$]. $\beta = 1$is the *p*-value of the *F*-stat. NoID refers to number of cross-sections. Number in parenthesis is White cross-section standard errors. * and ** indicate significant at 5% and 1% level respectively.

Table 9
Dynamic panel regression for U.S., Japan and Germany models from January 1995 to December 2009

A: U.S.	β	T-ratio	$\bar{R}^2$
OLS	0.988**	16.741	0.174
Bias-corrected-OLS	1.064**	12.754	0.173
FM-OLS	0.681**	8.091	0.157
Dynamic-OLS	0.641**	7.430	0.090
B: Japan			
OLS	0.748**	12.971	0.112
Bias-corrected-OLS	0.808**	9.264	0.111
FM-OLS	0.499**	5.673	0.100
Dynamic-OLS	0.566**	6.274	0.042

(*continued*)

Table 9 (*continued*)

C: Germany			
OLS	1.371**	19.055	0.215
Bias-corrected-OLS	1.490**	13.426	0.213
FM-OLS	0.797**	7.128	0.177
Dynamic-OLS	0.813**	7.086	0.126

Note: All regressions include unreported country-specific constants. The bias corrected *t*-statistics are reported in parentheses. ** denotes that the coefficient is significant at 1% level.

CONCLUSION

In this paper, we re-examine the well-known empirical puzzle of UIP using a sample of emerging economies. In particular, we focus on testing whether rejection of UIP is driven by the typically shorter horizons used in empirical studies.

The major finding of the paper is that the majority of emerging economies with more flexible exchange rate regimes clearly indicate that at longer maturity periods, the β coefficients of interest rate differentials for both time series and panel regressions are positive and getting closer to unity, as stated by UIP. The short-horizon finding confirms earlier results by Bansal and Dahlquist (2000), Frankel and Poonawala (2010), and Chinn and Meredith (2004; 2005), while the longer forecast horizon ($k = 12$) strengthens and expands those findings.

Complementing work on developed economies, this study has found a supportive ground to reconcile the theoretical-empirical puzzle of the UIP testing by adopting longer horizons for the exchange rate in emerging economies. This finding reveals that at the longer time horizon, the model has better econometric specification, more predictive power and less expectational error for exchange rate movements compared to the shorter time period, as explained by Chinn and Meredith (2005). Success or failure in testing UIP is sensitive to the selection of the prediction time horizon, *k*.

The findings can also be a signalling of well-integrated currency markets and a reliable guide to international investors as well as for the orderly conduct of monetary authorities. This signalling indicates that the benefit from international diversification borders may not be as high as previously understood, given the strong linkages between international monetary markets at a longer horizon. The evidence of co-integration implies that there is a common force, such as active arbitrage activity, which brings the exchange rate to "automatically adjust" in the long run. However, as pointed out by several authors, such as Francis et al.

(2002) and, Ferreira and Leon-Ledesma (2007) among others, co-integration does not rule out the possibility of arbitrage profit through diversification across markets in short-run terms, which may last for quite a while. Furthermore, it appears that domestic investors are becoming more aware of the economic interdependencies of international markets at a longer horizon by reacting to the developments in foreign markets and has increased capital mobility between markets in bringing world interest rates into line.

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NOTES

1. In practice, there is no sound basis for choosing other than the U.S. as a base country because 89% of exchange rate trading in the world uses the U.S. dollar.
2. We use the exchange rate regime definition provided by Reinhart and Rogoff (2004) in Appendix III.
3. However, it is not reported in Table 5 for brevity purposes and is available upon request from the author.

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